The Effect of Abortion Legalization on Teenage Out-of-Wedlock Childbearing in Future Cohorts

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Abstract

This paper examines the long-term impact of legalized abortion on teenage out-of-wedlock childbearing, which has been in constant decline since the early 1990s in the United States. Our argument is that, to the extent that it prevented unwanted births, legalized abortion could have reduced the likelihood of the teenage out-of-wedlock childbearing for the cohorts born after the legalization. We adopt a non-parametric approach that allows for a separate effect on Whites and African-Americans of the 1970 legalizations in the repeal states - California, New York, Washington, Alaska, and Hawaii – and the *Roe v. Wade* ruling in 1973. After controlling for the size of birth cohort, we find that for African-Americans, both changes lead to a long-term reduction in out-of-wedlock teenage childbearing. For Whites, there is no evidence supporting a long-term effect of the 1970 legalizations, but the cohorts born after *Roe v. Wade* in the non-repeal states show a reduction in teenage out-of-wedlock childbirth. Our findings are consistent with Levine et al. (1999), who find that the early legalization in the repeal states had a much stronger effect on the immediate fertility of Non-Whites than Whites. Also, our results show that legalized abortion can potentially account for at least 30 percent of the 45 percent decline in the teenage out-of-wedlock childbearing among 15-17 year olds for African-Americans and 35 percent of the 24 percent decline for Whites in the 1990s. Finally, since the fertility of African-Americans appears to be affected by legalized abortion earlier, our results suggest a potential reason for why teen out-of-wedlock childbearing for African-Americans started declining 3 years before than as it did for Whites.

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Key words: Abortion legalization; teenage out-of-wedlock childbearing; Roe v. Wade

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1. Introduction

Total teenage childbearing has been in a long-term decline since the late 1950s when it hit its highest level of 96 births per 1000 women aged 15-19, except for a brief but steep increase from the mid-1980s until the early 1990s. [1] The reasons for this decline are well documented; delayed marriage as well as reductions in the fertility of married teenagers being the most crucial explanations.[2] On the other hand, births to unmarried teenagers show a different pattern. Between 1960 and 1994, the non-marital birth rate per 1000 women aged 15-19 almost tripled, before going down after 1994.

It is interesting to note that while there are many studies that attempt to explain the reasons for the increase in the number of single mothers in general, as well as those of single teen mothers, studies examining the decline after 1994 are non-existent.[3] To our knowledge, only exception is a work by Levine (2001), though the outcome he studies is pregnancy risk of teenagers rather than their childbearing decisions. In that study, Levine covers a wide range of issues regarding decisions of teenagers on sexual activity, birth control use and childbearing. After extensively reviewing the previous literature on these issues, he examines the impact of "costs" associated with becoming pregnant and childbearing on pregnancy risk of teenagers. These costs include labor market conditions, the AIDS incidence, the generosity of the welfare system, and abortion restrictions in a teen's state of residence. His results suggest that more than a half of the decline in pregnancy risk among black, non-Hispanics that may have contributed the dramatic decline in teen fertility among this group cannot be explained by increases in these costs in the early to mid-1990s. Hence, we need to look at different avenues for a better and complete understanding of this crucial behavioral change.

This paper uses a non-parametric approach to investigate the legalization of abortion in the early 1970s in the United States as one of the possible explanations for this behavioral change.[4] Abortion was illegal under any circumstance in the United States

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[1] We discuss trends in teenage childbearing in the United States at length in Appendix 1.
[2] See Geronimous and Korenman (1992), Grogger and Bronars (1995, Hotz et al. (1997) to name a few studies for the negative consequences of giving birth in adolescence.
[3] We became aware after the completion of the first draft that Donahue, Grogger and Levitt (2002), in a work in progress, look at the same question using a different identification strategy. We give a detailed review of a very preliminary version of their study below.
[4] We are aware that there is an active debate in the United States regarding abortion. Our paper does not intend to take a position on this sensitive issue. Rather, it intends to extend the borders of the current

until the late 1960s.[5] This started to change between 1967 and 1970 when Colorado, North Carolina, California, Florida Georgia, Maryland, Arkansas, Kansas, New Mexico, Oregon, Delaware and Virginia made abortion legal for a limited number of circumstances. These circumstances included: i) to prevent the death or serious impairment of the physical and mental health of the mother; ii) the fetus would be born with a serious physical or mental defect; or iii) if the woman was raped or was subject to incest. The next step towards abortion legalization was taken in 1970 by Alaska, Hawaii, New York, and Washington when these states repealed their antiabortion laws, and by California when the California Supreme Court ruled that the state's abortion ban was unconstitutional. Thus in 1970, legal abortion became widely available in these states. In the sections follow, we will call these five states as *the repeal states* and the rest of the states as *the non-repeal states.* The final step towards abortion legalization was taken in 1973 in the U. S. Supreme Court ruling in *Roe vs. Wade*, which made abortion legal in the rest of the United States.

Legalized abortion could potentially account for part of the decline of teenage out-of-wedlock childbearing by changing the family environment in which teenagers were raised.[6]Also, by increasing the socioeconomic status of the mothers it could lead to an average lower non-marital fertility of their children. One set of studies suggests that cohorts born after legalized abortion experienced a significant decline in a number of adverse outcomes such as living in a single parent family, living in poverty, receiving welfare and infant mortality (see Gruber et al., 1999). Furthermore, Angrist and Evans (1999) find that African-American women who were exposed to abortion reforms experienced large reductions in teen fertility and teen out-of-wedlock fertility that appear to have led to increased schooling and employment rates. Another group of studies suggests that teens who grow up in a disadvantaged family are at greater risk of giving

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empirical research on out-of-wedlock teenage childbearing by suggesting a different avenue for investigation. Therefore it is a positive, not a normative, study.

[5] See Merz et al. (1995) for detailed history of state and federal laws concerning abortion in the 20th century, as well as before.

[6] As one might expect, the number of abortions performed in the US increased significantly during the 1970s and early 1980s before reaching its peak in 1983. Even though the data on the number of illegal abortions before the legalization are not available, if legal abortions simply replaced illegal abortions, then one would not see the increase in the number of abortion performed during the 1970s and early 1980s (DL, 2001).See Appendix 3 for abortion rates per 1000 live births in the repeal and non-repeal states during the 1970s and the early 1980s. .

birth out-of-wedlock (e.g. Moore et al., 1995; McLanahan and Sandefur, 1995). This paper, along with Donahue, Grogger, and Levitt (2002) (hereafter DGL), fills the gap between these two separate literatures and gives a different perspective on the issue of out-of-wedlock teenage childbearing trends observed in the United States.[7] Given this evidence, one might expect to see a relationship between the legalization of abortion in early the 1970's and the childbearing behavior in their adolescence for cohorts affected by the legalized abortion. Obviously we do not claim that abortion legalization is alone responsible for these trends in the 1990s, but we believe that it has the potential to account for a significant component of the trend.

The link we will examine is in the tradition of Donohue and Levitt (2001). In their seminal paper, DL investigate how legalized abortion influenced the criminal behavior of cohorts affected by the changes in abortion laws in the early 1970s. They found that this change accounts for most of the reduction in crime in the 1990s by reducing the number of unwanted births through selection. Since crime is mostly associated with males, the impact of the legalized abortion on crime works through aborted male fetuses after legalization. Abortion legalization would also be expected to have an impact on the outcomes of females such as teenage out-of-wedlock childbearing, on which we focus in this paper. We find that for African-Americans, both the 1970 legalizations in the repeal states – Alaska, California, Hawaii, New York, and Washington – and the *Roe v. Wade* decision lead to long-term reduction in out-of-wedlock teenage childbearing. For Whites, there is no evidence supporting a long-term effect of the 1970 legalizations, but the cohorts born after Roe v. Wade in the non-repeal states (the rest of the United States other than the repeal states) show a reduction in teenage out-of-wedlock childbirth. Our findings are consistent with Levine et al. (1999), who find that the early legalization in the repeal states had a much stronger effect on the immediate fertility of Non-Whites than Whites. Our model also passes a specification test indicating that it is indeed valid to take the legalization in the repeal states as exogenous. Finally, between 1994, the first year the increasing trend was reversed for Whites, and 2001 the decline in the out-of-wedlock birth rates among the 15-17 years olds was 24 percent for Whites. The same

[7] See also Ananat et al. (2006) for the selection effect of legalized abortion on several young adult outcomes.

decline was 45 percent for African-Americans between 1991 and 2001. Our results show that legalized abortion can potentially account for at least 30 percent of this percentage decline in the teenage out-of-wedlock childbearing among 15-17 years olds for African-Americans and 35 percent of this decline for Whites in the 1990s. Also since the fertility of African- Americans appears to be affected by legalized abortion earlier, our results suggest a potential reason for why teen out-of-wedlock childbearing for African-Americans started declining 3 years before than as it did for Whites.

The plan of the paper is as follows. In Section 2, we discuss further why there might be a potential long-term relationship between legalized abortion in the 1970s and the childbearing behavior of teens in the late 1980s and 1990s. We then review in Section 3 the literature on the potential reasons for giving birth out-of-wedlock in adolescence and on the impact of legalized abortion. Section 4 discusses the data. We present the econometric methodology used and our results in Section 5. After summarizing our findings and discussing their implications in Section 6, Section 7 concludes the paper.

2. A Potential Long-Term Effect of Legalized Abortion on Teenage Out-of-Wedlock Childbearing

Legalized abortion might have an impact on reducing per capita teenage out-of-wedlock childbearing through at least two channels. The first and more obvious channel is through facilitating the termination of unwanted pregnancies, which are more likely to be pregnancies of unmarried women.[8] The second channel, upon which this paper focuses, is an indirect one and expected to show its impact in the longer run. Potentially, females born after abortion legalization may have a lower likelihood of experiencing out-of-wedlock childbearing in adolescence. The first reason for this is that women who are more likely to have abortions, i.e. teenagers, unmarried women and the economically disadvantaged, are those potentially at risk of giving birth to children who end up having children out-of-wedlock themselves (Moore et al., 1995; McLanahan and Sandefour, 1995). Gruber et. al. (1999) document that the early life circumstances of the children on the margin of abortion are difficult along many dimensions: infant mortality, growing-up in a single-parent family and experiencing poverty. Adolescent females

[8] This first effect will drop out in the estimation strategies we will use below.

raised in these family environments are more likely to give birth out-of-wedlock. Also, legalized abortion provides a woman the opportunity to delay childbearing if current condition are suboptimal and allow her to give birth in the environment which she thinks is better and more nurturing for the development of her child (DL, 2001).[9] Finally, since the social and economic outcomes of black women were by pre-Roe reforms (Angrist and Evans, 1999), the increased socioeconomic status of the mothers may have lead to the average lower non-marital fertility of their children.

Before proceeding to our econometric analysis, we first present time series evidence on the teenage out-of-wedlock childbearing rates for the cohorts born between 1967 and 1982. Figures 1 and 2 respectively provide this for the total population and for African-Americans and Whites separately. The numbers in these figures can be interpreted as the cumulative number of births per 1000 women in the particular cohorts at the end of the teen years. For both the total population and African-Americans, teenage out-of-wedlock birth rates increase for the 1967-1973 birth cohorts (in the period 1982-1992). For the total population, this increase continues up until the 1976 birth cohort and then starts declining. The corresponding birth rates for African-Americans, however, show almost no change for the cohort born in 1974 and then start declining sharply for later cohorts. This time series evidence indicates that the trend for the non-marital birth rate of African-American teenagers experienced a break around the 1974 birth cohort, which is the first cohort affected by legalized abortion in the U.S among the non-repeal states.

3. Previous Literature on Teenage Fertility and the Impacts of Abortion Legalization

3.1 Previous Literature on the Effects of Abortion Legalization

There have been two *generations* of studies looking at the potential effects of legalized abortion. The *first generation* studies explored the immediate impact of abortion legalization on different outcomes such as women's fertility behavior (Baumann et al., 1977; Joyce and Mocan, 1990; Kramer, 1975; Levine et al., 1999 Quick, 1978;

[9] Note that if the male fetuses that were aborted were more likely to father a child out-of-wedlock, the reduced out-of-wedlock births in future cohorts could also be due to selection on males.

Sklar and Berkov, 1974; Tietze, 1973), schooling and labor market consequences (Angrist and Evans, 1999) and children's living conditions (Gruber et al., 1999). Further, Levine et al. (1999) looked at the effect of abortion legalization on births in the 1970s in the United States by using the variation in the timing of legalization across states in the early 1970's. They find that states legalizing abortion experienced a 5% decline in births relative to the rest of the country. The declines among teens, women over 35, and non-White women were even greater at 13%, 8% and 12% respectively. Angrist and Evans (1999) used the same variation to investigate the likely impact of legalization on education and labor market outcomes. Their results indicate that for White women, abortion reform did not appear to change schooling or labor market outcomes. On the other hand, African-American women who were exposed to abortion reforms experienced large reductions in teen fertility and teen out-of-wedlock fertility that appear to have led to increased schooling and employment rates.[10] Finally, Gruber et al. (1999) examine the effect of the increased availability of abortion after abortion legalization on the average living conditions of children. They found that cohorts born after legalized abortion experienced a significant decline in a number of adverse outcomes such as living in a single parent family or in poverty, receiving welfare and dying as an infant.

The *second generation* of studies started with the seminal paper by DL (2001). In this controversial paper, DL argued that legalized abortion contributed significantly to the reduction in crime in the 1990's. Since our paper is in the tradition of DL (2001), we will discuss their paper in more detail, starting with why one might expect to observe this relationship. In their paper, DL give two possible reasons. The first one is through reductions in cohort size resulting from abortion legalization. Smaller cohorts result in fewer young males in their primary crime years and thus less crime in the society. The more important reason is that children born after abortion legalization may have lower *per capita* crime rates since: (1) teenagers, unmarried women and economically disadvantaged women are those who are more likely to have abortion and also have the highest risk of giving birth to a child who could engage in criminal activity later in their lives; (2) abortion provides a woman the opportunity to delay childbearing if the current

[10] Note that since women with higher education are less likely to give birth to a child who will have an out-of-wedlock pregnancy, this effect could also contribute to lower non-marital pregnancy in the future cohorts.

conditions in her life are not suitable for raising a child. Their results suggest that this relationship between abortion legalization and crime exists because: (1) the five states that allowed abortion in 1970 experienced declines in crime earlier than the rest of the nation, which legalized abortion in 1973 with Supreme Court ruling on *Roe v. Wade*; (2) states with high abortion rates in the 1970s and 1980s experienced greater crime reductions in the 1990s; (3) in high abortion states, the number of arrests of "only" those born after abortion legalization fell relative to the low abortion states. This seminal article has initiated a renewed interest in the potential long-term impacts of legalized abortion. There have been papers examining the impact of abortion legalization on crime and other socioeconomic outcomes for other countries (Pop-Eleches; 2003, Sen, 2002) as well as on substance use (Charles and Stephens, 2002). Also, another group of studies has examined the robustness of the DL results using different identification strategies and data different data (Joyce 2004a, 2004b; Foote and Goetz, 2005).[11]

In a recent working paper, Ananat et al (2006) examine the selection impact of legalized abortion on several young adult outcomes such as living in poverty, being a single parent, receiving welfare, dropping-out of high school, graduating from college and being employed. They use the 2000 decennial census of the United States to measure their outcomes of interest for individuals born in a given state and year. Their main identification strategy is to estimate some innovative instrumental variables regressions using several variables potentially measuring the cost of abortion interacted with time dummies in the 1970s as their instruments. Their results indicate that lower costs of abortion led to improved outcomes for the cohorts affected by the law change through increasing the likelihood of college graduation, the lower use of welfare, and decreasing the likelihood of being single parent.

Finally and most importantly, DGL (2002), in a work in progress, investigate the effect of abortion legalization during the 1970s on teenage childbearing in the cohorts affected by the legalization.[12] They run a regression of the number of births on historical

[11] In fact, these studies argue that DL's results are sensitive to the identification strategy and the data used, so that their results are ambiguous at best. For DL's reply, see DL (2004) and (2006).
[12] Most of our work was carried out before we were aware of this paper. The only draft of the paper we were able access is Donahue et al. (2002), which is a very preliminary version. Therefore, all our discussions of this paper are based on this preliminary version. We thank Professor Levitt for pointing out their work, and that of Ananat et al. (2006), to us.

abortion rates in the period during which abortion was legalized controlling for state dummies and a number of control variables such as the current rate of abortions. At one point they use state dummies interacted with time dummies. They constrain the effect of abortions to be the same for Whites and African-Americans and for the 1970 legalization and *Roe v. Wade*. They find strong statistically significant results regarding the impact of historical abortion rates on teenage childbearing in the long-run.[13] Our work presented below differs from theirs in a number of ways. Most importantly, we estimate the effect of the legalizations on the birth rate directly. Also we allow these two legalizations to have different effects for African-Americans and Whites.[14] This distinction provides a strong test of the hypotheses of DL and DGL in the following sense. Levine (1999) has shown that the effect of the early legalization was much stronger for Non-Whites than Whites. Thus if the effect on out-of-wedlock births is through abortion, rather than through some other contemporaneous factor, we should see the African-American out-of-wedlock birth rates falling in the cohorts that were exposed to early legalization in the repeal states but among Whites we should not see this until we reach the White cohorts who were affected by *Roe v. Wade*. We do know that the out-of-wedlock birth rates for African-American teenagers fell three years before whites. However, this is only suggestive since only a fraction of African-American cohorts who were living in the repeal states were affected by the 1970 legalizations. In other words, it is important that the selection effect is seen for African-Americans in the states that were affected by early legalization, not just for African-Americans as a whole. In fact, this is exactly what we do see in the data, providing strong support for the DL interpretation.

One advantage of our approach of looking at the effect of the law directly is that it is possible that DGL might obtain inconsistent estimates using the abortion rate for two reasons. First, it is widely thought that reported abortion rates underestimate true abortion rates, which suggests that their estimates of the effect of the historical abortion rate might potentially be upward biased.[15] Secondly, abortion rates may be subject to classical measurement error. For example DL (2006) compares the abortion data collected by the

[13] Note that they also estimate their specifications for married teenagers.
[14] We use the birth rate as our dependent variable in state s in year t, while they use the number of births as the dependent variable. Using the birth rate has the advantage that it will not be affected by cohort size.
[15] Note, however, that their estimate of the coefficient times the number of abortions is consistent.

Allan Guttmacher Institute (AGI) with the data collected by Center for Disease Control (CDC) after controlling for some observable factors and obtains a correlation of only .396, which indicates a large measurement error in the data. This classical measurement error will lead to a downward bias for the coefficient on historical abortions.[16] Finally we are also able to use a specification test for our model, while DGL (2002) does not contain such a test.

In other aspects the differences between our work and DGL represent the standard differences between parametric and nonparametric work. Their estimates are more efficient if the restrictions behind their linear model are correct, but are inconsistent if the restrictions do not hold. Clearly there are arguments for both approaches.

3.2 Previous Literature on Teenage Childbearing

There is a rich literature on the causes, as well as the consequences, of teenage childbearing. Since our paper is connected to the former, we will only review studies investigating the causes related to our argument. [17] The first group of studies examine how individual factors, such as attitudes and expectations, risk taking behavior, school performance, and race influence teenage childbearing. These studies indicate that: (1) teens wanting a child or feeling ambivalent about having a child are more likely to experience adolescent childbearing (Zabin, 1994; Abrahamse et al., 1988; Hanson et al., 1989; Manlove, 1993); (2) educational expectations are negatively associated with the probability of adolescent childbearing (Moore et al., 1995b; Haggan and Wheaton, 1992; Sugland, 1992);[18] (3) teens engaging in risky activities are more likely to give birth (Hanson et al., 1989; Serbin et al., 1991); (4) girls who fall behind in school and girls with a more negative attitude toward school are more likely to have a non-marital teen birth (Moore et al., 1995b; Zabin, 1994; Plotnick and Butler, 1991); and (5) even though African-Americans are more likely to have non-marital births in their teen years (Abrahamse et al., 1988; Moore et al., 1995b), race is not statistically significant after controlling for crucial family, neighborhood and policy variables (Haveman and Wolfe, 1994).

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[16] Of course, this suggests that one could use the CDC number as an instrument for the AGI number to eliminate this second source of inconsistency.
[17] See Moore et al. (1995a) for an excellent review of studies on teenage childbearing.
[18] On the other hand, Plotnick and Butler (1991) could not find any association between expectations and teenage childbearing, or work attitudes and teenage childbearing.

Another group of studies analyzes the relationship between family background and teenage childbearing. The findings of these studies single out several important family variables associated with teenage childbearing. The major conclusions of these studies can be summarized as follows: (1) women raised in single parent families or who experienced parental marital separation are more likely to have out-of-wedlock births as teenagers (Moore et al., 1995b; Wu and Martinson, 1993; McLanahan and Sandefur, 1994; Kahn and Anderson, 1992; Haveman and Wolfe, 1994; Wu, 1994); (2) daughters of teen mothers are more likely to be teen mothers themselves (Kahn and Anderson, 1992; Manlove, 1993, 1995; Horrowitz et al., 1991); and (3) a higher maternal education is associated with a lower probability of teenage childbearing (Kahn and Anderson, 1992; Haveman et al., 1993).[19]

Last but not least, Levine (2001), in a study that is the most related to ours, examines the impact of "costs" associated with becoming pregnant and childbearing on pregnancy risk of teenagers using the multiple waves of the Youth Risk Behavior Survey (YRBS). These prices include labor market conditions, the AIDS incidence, the generosity of the welfare system, and abortion restrictions in a teen's state of residence. His results suggest that more than a half of the decline in pregnancy risk among black, non-Hispanics that may have contributed the dramatic decline in teen fertility among this group cannot be explained by increases in these costs in the early to mid-1990s. For whites, the impact of these changes in the costs seems even smaller.

4. Data

Data used in this paper comes from several sources. Birth data are obtained from the Vital Statistics of the United States.[20] They include the number of births to unmarried teenage women as well as the total live births to all teenagers by state, race and age for

[19] There are also a large numbers of studies in both the economics and sociology literatures on the effects of public policy variables, especially welfare generosity, on teenage out-of-wedlock childbearing. These studies find that the effect of larger welfare benefits on teen fertility is quite modest at best, and the inconsistencies of the research findings on this issue weaken even this conclusion. See Moffitt (1998) for excellent review of studies on the relationship between welfare policies and out-of-wedlock childbearing in general.

[20] One limitation of the Vital Statistics data is that we observe the states where teens give birth but not the teen's state of birth. Therefore we cannot account for selective migration occurring between the years when the teen was born and her childbearing age. Of course one can use decennial censuses to see how important this issue was for the birth cohorts we use. We will include this analysis in the future draft.

the period between 1981 and 2001.[21] Population data are provided by the National Cancer Institute (NCI). NCI has the county population estimates of the United States for the period covering 1969-2003. The estimates represent a modification by NCI of the annual time series of July 1 county population estimates by single year of age, sex and race that were originally produced by US Census Bureau.[22] Since we need state level population data, the county level estimates are aggregated to the state level. We use state level population data to obtain a variant of a *per capita* measure in order to isolate the long-term impact of abortion from the impact through *a smaller cohort size* due to abortion. As we have already stated, our measure will be the birth rate per 1000 women in a given demographic group. Finally, abortion data is provided by Alan Guttmacher Institute.

In order to isolate the different trends associated with African-American and White teenage out-of-wedlock childbearing, we carry out analysis of the two races separately, in addition to performing our analysis for the total population. Since Idaho, Maine, Montana, New Hampshire, North Dakota, South Dakota, Vermont, and Wyoming have a very low number of African-American teenagers and of births to African-American teenagers as a result, we drop these states from our data while performing our analysis for African-Americans.[23]

5. Econometric Methodology and Results

Our identification strategies exploit two important sets of legal changes regarding abortion laws that occurred in the early 1970s in the United States. The first set of legal changes took place in 1970 when Alaska, Hawaii, New York, and Washington repealed their antiabortion laws and the Supreme Court of California ruled in late 1969 that the state's law banning abortion was unconstitutional (DL, 2001). Second, we will take advantage of the historical ruling of the United States Supreme Court in *Roe v. Wade* in 1973 that lifted the ban on abortion in the rest of the country. We are not, of course, the first to exploit these changes. Indeed, as noted above, there is a rich literature on the

[21] We follow the conventional definition and classify women aged 15-19 as teenagers.
[22] For the full documentation of the modifications see http://seer.cancer.gov/popdata/
[23] In order to check whether our results for Whites are different due to the inclusion of these states in the regressions for Whites, we run our regression specifications using the same states that we used in our specifications for African-Americans. Our results did not change.

potential impacts of abortion legalization, and most of the recent papers in this literature used these changes in one way or another to obtain identification.

5.1 Abortion Legalization as a Natural Experiment: Threats to Identification

Finding a credible comparison group is very critical in a natural experiment exercise.[24] In this section, we summarize what can go wrong in identification if we consider abortion legalization as a natural experiment and how we will address those issues.

Suppose for simplicity we are interested in the long term impact of the 1970 legalization in the repeal states. Let the birth equation be given by

$$BR_{ijt} = \alpha_0 + \alpha_1 POST70_{ijt} + \gamma_i + \theta_j + \mu_i t + \delta_j t + \varepsilon_{ijt} \ . \qquad (1)$$

where BR_{ijt} is the birth rate of teenagers in state $\text{i} \, (\text{i} \in \text{repeal})$, age j and year t; $POST70_{ijt}$ is a dummy that is equal to one if the birth rate in state i, age j and year t belongs to teens who were born in the repeal states after abortion became legal; γ_i are state fixed-effects; θ_j are age fixed-effects; $\mu_i t$ capture state-specific trends; and $\delta_j t$ capture age-specific trends.

One possible approach is to choose i, j and t such that we compare teen non-marital birth rates of cohorts for the repeal states that were born before and after abortion became legal in the repeal states. This is the *Before-After* approach and can be written in the following way

$$\Delta BR_{ijt} = \alpha_1 \Delta POST70_{ijt} + \mu_i + \delta_j + \Delta\varepsilon_{it} \qquad (2)$$

where i = 1,…,5 ($i \in repeal$) and t = 1984,1986 and j=15. Thus those aged 15 in 1984 were not affected by legalized abortion in 1970 as they were born in 1969 while those aged 15 in 1986 were born in 1971 and affected by legalized abortion (see Figure 3). This

[24] See Meyer (1995) for excellent summary of potential treats to identification assumptions of natural experiments.

procedure will allow for treating γ_i and θ_j in (1) as fixed effects and δ_j as constant such that $\delta_j = \delta$ for all j. However, it will not allow one to capture the state-specific time trend coefficient μ_i. In other words μ_i acts like a fixed effect in first differences. Further, even if $\mu_i = \mu$ for all i one would not be able to separately identify α_1 from μ. The latter problem can be addressed by, for example, using equation (2) and setting:

i = 1,...,5 (i $\in$ repeal); i' = 6,...,51 (i' $\in$ nonrepeal); t =1984, 1986; and j = 15 as long as $\mu_i = \mu_{i'} = \mu$ for all i and i' so that

$$\Delta\overline{BR}_{(r)jt} - \Delta\overline{BR}_{(nr)jt} = \alpha_1 \Delta POST70_{(r)jt} + (\Delta\bar{\varepsilon}_{(r)jt} - \Delta\bar{\varepsilon}_{(nr)jt}), \tag{3a}$$

$$\Delta\overline{BR}_{(r)jt} = \frac{\sum_{i\in repeal} \Delta BR_{ijt}}{N_{repeal}} \text{ and } \Delta\overline{BR}_{(nr)jt} = \frac{\sum_{i'\in nonrepeal} \Delta BR_{i'jt}}{N_{nonrepeal}}, \tag{3b}$$

where N_{repeal} and $N_{nonrepeal}$ are the numbers of the repeal and non-repeal states respectively. Notice that estimating above regression using the difference of first differences in non-marital teenage birth rates between the repeal and non-repeal states eliminates μ. This will be the first estimator we will use and refer to it as the *difference-in-difference estimator (DD).* Of course in practice it is inefficient to use only those aged 15 years old, and I will also use 16, 17, 18 and 19 year-olds for our *DD* estimation.

However, if $\mu_i \neq \mu_{i'} \neq \mu$ for all i and i', one has the problem so that equation (3) will be as follows[25]

$$\Delta\overline{BR}_{(r)jt} - \Delta\overline{BR}_{(nr)jt} = \alpha_1 \Delta POST70_{ijt} + (\bar{\mu}_{(r)} - \bar{\mu}_{(nr)}) + (\Delta\bar{\varepsilon}_{(r)jt} - \Delta\bar{\varepsilon}_{(nr)jt}) \tag{4a}$$

$$\bar{\mu}_{(r)} = \frac{\sum_{i\in repeal} \mu_i}{N_{repeal}} \text{ and } \bar{\mu}_{(nr)} = \frac{\sum_{i'\in nonrepeal} \mu_{i'}}{N_{nonrepeal}} \tag{4b}$$

To address this problem, consider (4a) for another age group j' ($j' \neq j$)

[25] Note that $\Delta POST70_{i'jt}$ is zero since abortion was still illegal in the non-repeal states in 1971.

$$\Delta\overline{BR}_{(r)j't} - \Delta\overline{BR}_{(nr)j't} = \alpha_1 \Delta POST70_{(r)j't} + (\overline{\mu}_{(r)} - \overline{\mu}_{(nr)}) + (\Delta\overline{\varepsilon}_{(r)j't} - \Delta\overline{\varepsilon}_{(nr)j't}) \qquad (5)$$

For example we could look at the first differences of 15 and 17 year-olds between 1986-1984 (see Figure 4), and subtract (4a) from (5). Taking this extra difference will eliminate state-specific time trends in the data since

$$(\Delta\overline{BR}_{(r)jt} - \Delta\overline{BR}_{(nr)jt}) - (\Delta\overline{BR}_{(r)j't} - \Delta\overline{BR}_{(nr)j't}) = \alpha_1 \Delta POST70_{(r)jt} + ((\Delta\overline{\varepsilon}_{(r)jt} - \Delta\overline{\varepsilon}_{(nr)jt}) \\ -(\Delta\overline{\varepsilon}_{(r)j't} - \Delta\overline{\varepsilon}_{(nr)j't})) \qquad (6)$$

This is the *Difference-in-Difference-in-Difference (DDD)* estimator and in this case α_1 identifies the long-term impact of the 1970 legalizations in the repeal states allowing for state specific trends. [26] In the sections follow, we will use the more conservative *DD* and *DDD* approaches and disregard the *Before-After* estimation.

5.2 Cross-State DD Estimation: A Comparison of Early Legalization States with the Rest of the United States

In this part of the paper, we first define the birth cohort years 1967-1970 as the pre-legalization years and 1971-1973 as the post-legalization years in the repeal states. Teenagers who were born in the pre-legalization years, whether they were born in the early legalization states or in the rest of the U. S., would not be affected by the changes in the abortion law that took place in 1970. On the other hand, for the cohorts born between 1971 and 1973 exposure to the legalized abortion will differ among teenagers; if they were born in the repeal states they would be affected, otherwise they would not. Therefore, our DD will capture the change in birth rates of single teenagers between pre-

[26] Notice that $\alpha_1 \Delta POST70_{(r)j't}$ is zero since age group j' in the repeal states did not affected by the 1970 legalizations since they were born before 1970.

and post-legalization years in the repeal states relative to the changes in the non-repeal states. If we want to put DD in a regression framework, it can be written as follows[27]

$$\ln BR_{ijt} = \gamma_0 + \gamma_1 (REPEAL_i \times C7173_{jt}) + a_j \bullet s_j + a_j \bullet y_t + u_{ijt}, \qquad (7)$$

where $\ln BR_{ijt}$ is the logarithm of birth rates per thousand teen of age j, in state i and year t; $REPEAL_i$ is the dummy variable equal to one if state i is among the early legalizing states and $C7173_{jt}$ is the 1971-1973 birth cohort dummy for year t and age j.[28] Further, we also control for age-state ($a_j \bullet s_j$) as well as age-year ($a_j \bullet y_t$) effects. In the above regression γ_1 is our parameter of interest, DD, and can be interpreted as the proportionate change in birth rates between the 1971-1973 and 1967-1970 birth cohorts in the repeal states relative to the rest of the United States. We expect this coefficient to be negative since our conjecture is that the teenagers who were born in the repeal states between 1971 and 1973 are less likely to have a non-marital birth relative to the teenagers who were born in the non-repeal states during the same time period. Our assumption here is that there are no state-specific time trends so that the non-repeal states form a credible control group to examine the impact of legalized abortion for the repeal states.

Table 1 presents DD estimation results. All regressions are carried out by weighted least squares, where the square root of the state teenage population for the corresponding race for a given year is the weight for a given observation. We estimate the impact separately for African Americans and Whites. We believe this provides a strong test of the idea that legalized abortion affected teenage out-of-wedlock childbearing. Levine et al. (1999) have shown that the early legalization in the repeal states had a stronger effect on the immediate fertility of Non-Whites than Whites in the early 1970s,

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[27] Notice that this regression is in levels rather than in differences. We also use more than one age group and include the 1972 and 1973 birth cohorts compare to a simple example of *DD* we described in the previous section. Therefore it is more general compare to equation (3a) but γ_1 still identifies *DD* estimator.

[28] For example, consider the observations for 1988. For this year, $C7173_{88}$ will be equal to 1 for 15, 16, and 17 years old women but zero for women aged 18 and 19. Note also that this is identical to $POST70_{ijt}$ dummy in the previous section.

so it should be the case that the effect of the early legalization was potentially stronger for African Americans than Whites. [29]

The first panel of the table presents the estimates for the total teenage population. Panels 2 and 3 show our estimates for the African-American and White teenage populations respectively. The first column of all the panels shows the results of our baseline specification. While for the total population and Whites, coefficients on $REPEAL_i \times C7173_{jt}$ are not significantly different from zero, the coefficient is much larger and statistically significant for African-Americans. Specifically, teenage out-of-wedlock childbearing declined about 10 percent for African-Americans in the repeal states relative to African-Americans in the non-repeal states for the 1971-1973 birth cohorts.

In order to check the credibility of the teenagers in the non-repeal states as a comparison group, we carry out several specification tests. First, if the teenagers in the non-repeal states are to be a credible comparison group, then trends in non-marital teenage childbearing for the repeal and non-repeal states must be comparable before the repeal states legalized abortion. To test this conjecture, we consider the following regression

$$\ln BR_{ijt} = \gamma_0 + \gamma_1(REPEAL_i \times C7173_{jt}) + \gamma_2(REPEAL_i \times C70_{jt}) + a_j \bullet s_i + a_j \bullet y_t + u_{ijt} \qquad (8)$$

where $C70_{it}$ is the 1970 birth cohort dummy for year t and age j, so that the interaction term is equal to one only for teens from repeal states who were born in 1970. The coefficient γ_2 captures the proportionate change in birth rates between the 1970 and 1967-1969 birth cohorts in the early legalizing states relative to the rest of the United States. If the trend in the out-of-wedlock teenage childbearing started diverging with the 1971 birth cohort, then this coefficient should be close to zero and not statistically significant. All other regression parameters are the same as in equation (7). This is the specification test that we refer to in the introduction.

[29] Here our assumption is the selection impact of legalized abortion for a birth aborted by African-Americans and Whites are the same so that lower fertility rates because of legalized abortion in the part of African-Americans in the 1970s should reflect a greater selection and a greater long-run impact.

Second, we expand our sample by adding data on childbearing behavior of teens born between 1974 and 1979, and add three extra interaction terms to our regression equation (8)

$$\begin{aligned}\ln BR_{ijt} = \gamma_0 + \gamma_1(REPEAL_i \times C7173_{jt}) + \gamma_2(REPEAL_i \times C70_{jt}) \\ +\gamma_3(REPEAL_i \times C74_{jt}) + \gamma_4(REPEAL_i \times C7576_{jt}) \\ +\gamma_5(REPEAL_i \times C7779_{jt}) + a_j \bullet s_i + a_j \bullet y_t + u^*_{ijt}\end{aligned} \tag{9}$$

where $C74_{jt}$ $C7576_{jt}$ and $C7779_{jt}$ are the birth cohort dummies for teens of age j in year t who were born in 1974, 1975-1976 and 1977-1979 respectively. Therefore, the coefficients γ_3, γ_4, and γ_5 capture how teenage childbearing in the non-repeal states affected by *Roe v. Wade* differs from that for the repeal states for the cohorts who were born in the years following *Roe v. Wade*. The coefficients γ_3, γ_4, and γ_5 should not be significantly different from zero if (1) legalized abortion was the only factor that accounted for the potentially diverging trends in the teenage childbearing for cohorts born in the 1971-1973 period between repeal and non-repeal states so that there was no state-specific time trends affecting birth rates; (2) both the 1970 legalizations in the repeal states and *Roe v. Wade* had the same impact on the out-of-wedlock childbearing of the birth cohorts affected by these law changes towards abortion. Thus, this is the specification test of the assumption underlying *DD*.

In the second column of Table 1 we present the results from these tests. The coefficients on $REPEAL_i \times C70_{jt}$ are not statistically different from zero for any of the groups. Thus our model passes a specification test which indicates that it is reasonable to treat the 1970 legalization in the repeal states as exogenous. The coefficients on $REPEAL_i \times C74_{jt}$, $REPEAL_i \times C7576_{jt}$, and $REPEAL_i \times C7779_{jt}$ are not statistically different from zero as the coefficient for the total population. For African-Americans, on the other hand, they are negative and as large as the coefficient on $REPEAL_i \times C7173_{jt}$. This may seem to be a surprising result since one might expect catching-up on the part of non-repeal states after *Roe v. Wade* legalized abortion in the rest of the country. But Levine et al. (1999) show that while the birth rates of Whites living the non-repeal and repeal states had converged after *Roe v. Wade*, for African-Americans this difference stayed about the same, with the rebound in births in the non-repeal states being small. For Whites who were born between 1970-1973, even though it seems that are no statistically significant differences in the change of teenage childbearing between repeal and non-repeal states, this difference becomes positive and statistically significant for cohorts born between 1974-1976. However, it becomes statistically insignificant and close to zero for the 1977-1979 birth cohorts.

We can summarize our findings in this section as follows. First, while the 1971-1973 African-American birth cohorts in the repeal states were 10 percent less likely to have non-marital childbearing in their teen years relative to their peers in the non-repeal states, we do not find any difference for Whites. Second, neither for African-Americans nor for Whites, the trends in the out-of-wedlock teenage childbearing differ between the repeal states and the non-repeal states for the birth cohorts that were not affected by the 1970 legalizations. Finally, the non-marital teenage childbearing of African-American cohorts born after *Roe v. Wade* did not converge between the repeal states and the non-repeal states for the either or both of the following reasons: (1) State-specific trends confounds with our *DD* estimates; (2) the 1970 legalizations in the repeal states and *Roe v. Wade* had different impacts on the out-of-wedlock childbearing of the birth cohorts affected by these law changes towards abortion.

To address these issues, we will first estimate the impact of the 1970 legalizations in the repeal states using a *DDD* strategy, which controls for possible state-specific trends. We will then estimate the impact of *Roe v. Wade* separately using a similar *DDD* strategy to compare its impact in the non-repeal states to the impact of the 1970 legalizations in the repeal states.

5.3 Difference-in-Difference-in-Difference (DDD) Estimation of the Impact of 1970 Legalizations

This strategy is also used by Joyce (2004) to investigate the impact of legalized abortion on the crime rate in the United States.[30] The identifying assumption is that there

[30] Our DDD specification is comparable to one of DGL's specifications in which they use state*time interactions .

is no distinct contemporaneous shock that affects the non-marital birth rates of teenagers affected by legalized abortion in the repeal states relative to those of teens who were not in the same state and year.

In Section 5.1 we explained how this estimator can be calculated for 15 year-olds who were born in 1971 and thus affected by the 1970 legalizations using 17 year olds as a comparison group. We can also compute the same DDD measure for the 16 and 17 years olds who were born in 1971 in the repeal states by using the 18 and 19 years olds as their respective comparison groups. Our final DDD estimate will be the average of these three measures. Note that using these particular cohorts and ages we observe the birth rates for 15 and 17 year-olds, 16 and 18 year-olds, and 17 and 19 year-olds in the same years so that both age groups exposed to same period effects (1984-1988).

This estimator could also be illustrated as follows:

$$DDD = \underbrace{(\Delta\overline{BR}_{(r)jt} - \Delta\overline{BR}_{(nr)jt})}_{DD_{affected}} - \underbrace{((\Delta\overline{BR}_{(r)j't} - \Delta\overline{BR}_{(nr)j't}))}_{DD_{unaffected}} \qquad (10)$$

t year and, j and j' are age subscripts such that

$$\{(j;j';t)\} = \{(15;17;1984),(16;18;1985),(15;17;1986),(17;19;1986),(16;18;1987),(17;19;1988)\}$$

are the complete account of years and ages we use in the estimation. Further, $\Delta\overline{BR}_{ijt}$ and $\Delta\overline{BR}_{ij't}$ are the first differences in the mean non-marital birth rates in the repeal states for age groups j and j' as described in (3b). $\Delta\overline{BR}_{i'jt}$ and $\Delta\overline{BR}_{i'j't}$ are the same first differences in the non-repeal states for age groups j and j'.

The stacked version of this *DDD* estimation can be estimated in the logarithmic form as running the following regression:

$$\begin{aligned} \ln BR_{ijt} &= \delta_0 + \delta_1 AGEAFFECTED_j + \delta_2(AGEAFFECTED_j \times POST70_t) \\ &+\delta_3(AGEAFFECTED_j \times REPEAL_i) + \delta_4(REPEAL_i \times POST70_t) \qquad (11) \\ &+\delta_5(AGEAFFECTED_j \times REPEAL_i \times POST70_t) + s_i + y_t + e_{ijt}\,. \end{aligned}$$

where $AGEAFFECTED_j$ is a dummy variable taking value of one for age group affected by legalized abortion and zero otherwise; $REPEAL_i$ is the dummy variable equal to one if state i is among the early legalizing states; and $POST70_t$ is a dummy being one for the post legalization years. Further, s_i and y_t are state and year dummies respectively.[31] In equation (10) the coefficient of the second order interaction (δ_5) captures the DDD estimate while other terms in (10) can be thought of independent variables controlling for state, year, and state-year effects.[32]

The first panel of Table 2 presents the *DDD* results for the total population, African-Americans and Whites respectively. They show that while there is no statistically significant reduction in non-marital teenage childbearing for Whites there is a nearly 5 percent statistically significant decline for the African-American cohorts born after abortion became legal in the repeal states in 1970.[33] While these estimates are consistent with our *DD* estimate for Whites in the previous section, the effect for African-Americans declines to about 5 percent from our previous estimate of 10 percent but is still statistically significant.

5.4 Difference-in-Difference-in-Difference (DDD) Estimation of the Impact of Roe v. Wade

In 1973 the U.S. Supreme Court ruling in *Roe v. Wade* legalized abortion in the rest of the U.S. One may think of this important decision as a kind of treatment reversal. Although women living in the repeal states had access to abortion after *Roe v. Wade* and women in these states had a head start in terms of abortion access, one would still expect to see an impact for the non-repeal states relative to the repeal states considering the extent of the change in non-repeal states. The figures that we presented before suggest that while for Whites there is indeed a convergence in non-marital childbearing rates

[31] Notice that in (13) we do not include state-year dummies since the other variables included in the regression controls for state specific trends. See Appendix 2.

[32] For those unfamiliar with *DDD*, it may not be obvious that δ_5 is the parameter of interest. This is shown in Appendix 2.

[33] Clearly, the *DDD* estimate for the total population is a weighted average of *DDDs* for different racial groups. Since there appears to be a different impact for African-American and White treatment groups, we only will discuss our findings for these groups and ignore the estimates for the total population.

between repeal and non-repeal states, this is not so for African-Americans. But these trends might conceal the real impact of *Roe v. Wade*, since there are potentially many factors that might affect the childbearing decision of the teens, and thus it possible to confound the shock affecting all the teenagers in a particular state preventing teenage birth rates from converging to their pre-legalization trends. In order to isolate such a shock, we again consider the *DDD* estimation strategy to investigate the long term impact of *Roe v. Wade*. As before the *DDD* estimate will be the average of DDD measures of 15, 16 and 17 years olds and can be written as follows

$$DDD_{Roe} = \underbrace{(\Delta\overline{BR}_{(r)jt} - \Delta\overline{BR}_{(nr)jt})}_{DD_{affected}} - \underbrace{((\Delta\overline{BR}_{(r)j't} - \Delta\overline{BR}_{(nr)j't}))}_{DD_{unaffected}} \qquad (12)$$

where t is year, and j and j' are age subscripts such that

$$\{(j;j';t)\} = \{(15;17;1987),(16;18;1988),(15;17;1989),(17;19;1989),(16;18;1990),(17;19;1991)\}$$

are the complete age and years used in the estimation. The stacked version of this *DDD* estimation can be estimated in the logarithmic form as running the following regression

$$\begin{aligned} \ln BR_{ijt} = {} & \beta_0 + \beta_1 AGEAFFECTED_j + \beta_2(AGEAFFECTED_j \times POST73_t) \\ & + \beta_3(AGEAFFECTED_j \times NONREPEAL_i) + \beta_4(NONREPEAL_i \times POST73_t) \\ & + \beta_5(AGEAFFECTED_j \times NONREPEAL_i \times POST73_t) + s_j + y_t + \omega_{ijt} \end{aligned} \qquad (13)$$

$AGEAFFECTED_j$ is again a dummy variable taking a value of one for age group exposed to legalized abortion after *Roe v. Wade* and of zero otherwise; $NONREPEAL_i$ is a dummy variable equal to one if state i legalized abortion after *Roe v. Wade* ; $POST73_t$ is a dummy being one for the post-Roe years; s_i and y_t are state and year dummies respectively.

In equation (13) the coefficient β_5 provides the *DDD* estimate. Since abortion was legal in the repeal states before *Roe v. Wade* and the abortion rates in those states were still increasing during this time period, our coefficient potentially will be biased downward. In other words, *Roe v. Wade* can be thought of as a treatment that should have only affected the cohorts that were born in the non-repeal states after *Roe v. Wade*.

However, in the repeal states the impact was the ongoing impact of the 1970 legalizations but not *Roe v. Wade* itself. Hence, our estimate is potentially a lower bound of the true effect of legalized abortion for teens born after *Roe v. Wade* in the non-repeal states.

The second panel of Table 2 presents *DDD* estimates for the long-term impact of *Roe v. Wade* in the non-repeal states. As discussed before, abortion was legal in the repeal states before the *Roe v. Wade*, and the abortion rates in those states were still increasing during this time period. Hence, our estimates will be biased downward potentially. Surprisingly, our coefficients are large and statistically significant for both African-Americans and Whites. There are 13 percent and 9 percent declines respectively in the teenage out-of-wedlock birth rates for African-Americans and Whites affected by *Roe v. Wade*. Considering that these relatively large estimates are potentially lower bounds for the true effects of *Roe v. Wade*, one might conclude that it had a substantial impact on the teenage out-of-wedlock childbearing behavior of both African-Americans and Whites. [34]

6. Discussion

Our results in the previous section can be summarized as follows. First, abortion in 1970 had a significant impact on the childbearing behavior of African-American teens who were affected by the law change. However, the estimated impact declined by half when we control for state-specific time trend and we find no effect for Whites living in the repeal states. Second, *Roe v. Wade* affected the birth rates of both groups living in the non-repeal states but the effect was larger for African-Americans. Third, the impact of *Roe v. Wade* in the non-repeal states was larger than that of the earlier legalization in the repeal states for both races.[35]

Our findings have several noteworthy implications. Between 1994, the first year the increasing trend was reversed for Whites, and 2001 the decline in the out-of-wedlock

[34] To check the robustness of our results, we also used all teenage births in another specification and obtained very similar estimates.

[35] One of the potential explanations for this finding is that since we do not account for the residents of the non-repeal states that commuted to the repeal states to obtain an abortion between 1970 and 1973, our estimates for the impact of the 1970 legalization might be downward biased. In order to deal with this problem, in another specification we dropped the non-repeal states, which are adjacent to the repeal states, from our sample. Our estimates using this smaller sample are not significantly different from the results presented in Table 2.

birth rates among the 15-17 years olds was 24 percent for Whites. The same decline was 45 percent for African-Americans between 1991 and 2001. Our the *DDD* estimates of the impact of *Roe v. Wade* that control for state-specific time trends show that legalized abortion can potentially account for at least about 30 percent of this decline in the teenage out-of-wedlock childbearing among 15-17 years olds for African-Americans and about 35 percent of this decline for Whites in the 1990s.[36] Finally, since the fertility of African-Americans appears to be affected by legalized abortion earlier, our results suggest a potential reason for why teen out-of-wedlock childbearing for African-Americans started declining 3 years before than as it did for Whites.

7. Conclusion

In this paper, we attempt to bridge the gap between two different literatures in order to give a different perspective on and a better understanding of the one of the important trend changes in the early 1990's. Teenage out-of-wedlock childbearing has been in decline for more than a decade now, and studies - with the exception of Levin (2001) and DGL - (2002), considering why this behavioral change occurred are non-existent in the academic literature.[37] First, our *DD* estimation provides a specification test of whether it is reasonable to treat the early legalization in the repeal states as exogenous, and we find that the model passes this test. Our findings show that for African-Americans both the 1970 legalization in repeal states and *Roe v. Wade* lead to long-term reduction in out-of-wedlock teenage childbearing. For Whites, there is no evidence supporting a long-term effect of the 1970 legalizations, but the cohorts born after Roe v. Wade in the non-repeal states show a reduction in teenage out-of-wedlock childbirth. Our results also offer a potential explanation for why teenage out-of-wedlock childbearing for African-Americans started

[36] We should mention that the 1970s was the time period when the access of unmarried women to oral contraceptives was also increasing throughout the U.S. Therefore, to the extent that oral contraceptives prevented unwanted childbearing of unmarried women in the 1970s, one can argue that the effect we identify is due to oral contraceptives, at least a part of it, but not legalized abortion. While we believe that the increased availability of oral contraceptives in the 1970s potentially had a long-term impact on teenage out-of-wedlock childbearing similar to legalized abortion, without any radical change in the usage of pills around the time of legalizations in 1970 in the repeal states and in 1973 in the non-repeal states, our results will not be affected from this confounding impact significantly.

[37] Darroch and Singh (1999) document changes in teen sexual behavior, but do not try to explain why the behavior changed.

declining three years before that of Whites. Our the *DDD* estimates of *Roe v. Wade* that control for state-specific time trends show that legalized abortion can potentially account for at least about 30 percent of this percentage decline in the teenage out-of-wedlock childbearing among 15-17 years olds for African-Americans and about 35 percent of this decline for Whites in the 1990s.

References

Abrahamse, A. F., P.A. Morrison and L.J. Waite. 1988. "Teenagers Willing to Consider Single Parenthood: Who is at Greatest Risk?" *Family Planning Perspectives* 20: 13-18.

Alan Guttmacher Institute. 2005. *Alan Guttmacher State Abortion Trends*, Unpublished Abortion Data, New York, NY: Alan Guttmacher Institute.

Ananat, E. O., J. Guber, P. B. Levine and D. Staiger. 2006. "Abortion and Selection" NBER Working Paper No. 12150.

Angrist, J. and W. Evans. 1999. "Schooling and Labor Market Consequences of the 1970 State Abortion Reforms." *Research in Labor Economics* 18, S. Polachek ed., Greenwich: JAI Press.

Bachrach, C. and K. Carver. 1992. "Outcomes of Early Childbearing: An Appraisal of Recent Evidence." Summary of a Conference. National Institute of Child Health and Human Development.

Bronars, S. and J. Grogger. 1995. "The Economic Consequences of Unwed Motherhood: Using Twin Births as a Natural Experiment." *American Economic Review*, 84(5): 1141-1156.

Charles, K.K., and M. J. Stephens. 2002. "Abortion Legalization and Adolescent Substance Abuse." NBER Working Paper No. 8571.

Census Bureau. *Vital Statistics of the United States: Natality*, Washington, DC: U. S. Census Bureau.

Darroch, J.E. and S. Singh. 1999. "Why Teenage Pregnancy Declining? The Roles for Abstinence, Sexual Activity and Contraceptive Use." Alan Guttmacher Institute Occasional Report No. 1 New York, NY: Alan Guttmacher Institute.

Donohue, J. and S. Levitt. 2001. "The Impact of Legalized Abortion on Crime." *Quarterly Journal of Economics*, 116(2):379-420.

Donohue, J. and S. Levitt. 2004. "Further Evidence that Legalized Abortion Lowered Crime: A Reply to Joyce." *Journal of Human Resources*, 2004, 39(1): 29-49.

Donohue, J. and S. Levitt. 2006. "Measurement Error, Legalized Abortion, the Decline in Crime: A Response to Foote and Goetz(2005)." NBER Working Paper No. 11987.

Donohue, J., S. Levitt and J. Groger. (2002). "The Impact of Legalized Abortion on Teen Childbearing." ABF Working Paper No. 2112.

Foote, C. and C. Goetz. 2005. “Testing Economic Hypothesis with State-Level Data: A Comment on Donohue and Levitt (2001).” Federal Reserve Bank of Boston Working Paper 05-15.

Geronimus, A. and S. Korenman.1992. “The Socioeconomic Consequences of Teen Childbearing Reconsidered.” *Quarterly Journal of Economics*, 107: 1187-1214.

Gruber, J., P. Levine, and D. Staiger. 1999. “Abortion and Child Living Circumstances: Who is the ‘Marginal Child’?” *Quarterly Journal of Economics*, 114(1): 263-291.

Haggan, J. and B. Wheaton. 1993. “The Search for Adolescent Role Exits and the Transition to Adulthood.” *Social Forces* 71(4): 955-980.

Hanson, S.L., D.R. Morrison and A.L. Ginsberg. 1989. “The Antecedents of Teenage Fatherhood.” *Demography* 26: 579-596.

Haveman, R. B. Wolfe, and C. B. An. 1993. “Teen Out-of-Wedlock Births and Welfare Receipt: The Role of Childhood Events and Economic Circumstances.” *Review of Economics and Statistics* 75(2): 195-208.

Horowitz, S.M., L.V. Klerman, H.S. Kuo and J.F. Jekel. 1991. “Intergenerational Transmission of School Age Parenthood.” *Family Planning Perspectives* 23(4):168-172.

Hotz, V. J., S. McElroy and S. Sanders. 1997. “The Costs and Consequences of Teenage Childbearing for the Mothers and the Government.” in *Kids Having Kids: The Economics Costs and Social Consequences of Teen Pregnancy* R. Maynard ed., Washington, DC: Urban Institute Press.

Joyce, T. J. 2003. “Did Legalized Abortion Lower Crime?” *Journal of Human Resources* 38(1):1-37.

Joyce, T.J. 2004. “Further Tests of Abortion and Crime.” NBER Working Paper No. 10564.

Joyce, T.J. and N.H. Mocan. 1990. “The Impact of Legalized Abortion on Adolescent Childbearing in New York City.” *American Journal of Public Health* 80(3): 273-278.

Kahn, J.R. and K.E. Anderson. 1992. “Intergenerational Patterns of Teenage Fertility.” *Demography* 29(1):39-57.

Kramer, M.J. 1975. “Legal Abortion among New York City Residents: An Analysis According to Socioeconomic and Demographic Characteristics.” *Family Planning Perspectives* 7: 128-137.

Levine, D. I. and G. Painter. 2003. "The Schooling Costs of Teenage Out-of-Wedlock Childbearing: Analysis with a Within-School Propensity Score Matching Estimator." *Review of Economics and Statistics* 85(4): 884 – 900.

Levine, P.B. 2001. "The Sexual Activity and Birth Control Use of American Teenagers." In *An Economic Analysis of Risky Behavior among Youths*, Jonathan Gruber ed. Chicago: Chicago University Press.

Levine, P.B., D. Staiger, T. Kane, and D.J. Zimmerman. 1999. "Roe vs. Wade and American Fertility." *American Journal of Public Health* 89(2): 199-203.

Manlove, J. 1993. "The Social Reproduction of Teenage Motherhood in Great Britain." Unpublished Doctoral Dissertation, Duke University: Durham, NC

Manlove, J. 1995."Breaking the Cycle of Disadvantage: Ties between Educational Attainments, Dropping Out and School-Age Motherhood." Presented at the *Annual Meeting of the American Educational Research Association*, San Francisco, CA.

McLanahan, S. and G. Sandefur. 1995. *Growing Up with a Single Parent: What Hurts, What Helps.* Cambridge, MA: Harvard University Press.

Merz, J.F., C.A. Jackson, and J. A. Klerman. 1995. "A Review of Abortion Policy; Legality, Medicaid Funding, and Parental Involvement, 1967-1994." *Women's Rights Law Reporter* 17: 1-61.

Meyer, B. D. 1995. "Natural and Quasi- Experiments in Economics." *Journal of Business and Economic Statistics* 13: 151-162.

Moffitt, R. 1998. "The Effect of Welfare on Marriage and Fertility." in *Welfare, the Family and Reproductive Behavior*, Robert Moffitt ed. Washington DC.: National Academy Press.

Moore, K.A., B.C. Miller, D. Glei and D.R. Morrison. 1995a. *Adolescent Sex, Contraception, and Childbearing: A Review of Recent Research.* Washington, D.C.: Child Trends, Inc.

Moore, K.A., D. Glei, D.R. Morrison. and J. Manlove. 1995b. "Non-marital Childbearing: Family, Peer, and School Influences." Unpublished Manuscript, Washington, DC: Child Trends, Inc.

Quick, J.D. 1978. "Liberalized Abortion in Oregon: Effects on Fertility, Prematurity, Fetal Death, and Infant Death." *American Journal of Public Health* 68: 1003-1008.

Plotnick, R.D. and S.S. Butler. 1991. "Attitudes and Adolescent Non-marital Childbearing: Evidence from the National Longitudinal Survey of Youth." *Journal of Adolescent Research* 6: 470-492.

Pop-Eleches, Christian. 2003. “The Impact of an Abortion Ban on Socio-Economic Outcomes of Children: Evidence from Romania.” Harvard University, Department of Economics, Unpublished.

Sen, A. 2002. “Does Increased Abortion Lead to Lower Crime? Evaluating the Relationship between Crime, Abortion and Fertility.” University of Waterloo, Department of Economics. Unpublished.

Serbin, L.A., P.L. Peters, V.J. McAffer and A.E. Schwartzman. 1991. “ Childhood Aggression and Withdrawal as Predictors of Adolescent Pregnancy, Early Parenthood, and the Environmental Risk for the Next Generation.” *Canadian Journal of Behaviourial Science* 23(3): 318-331.

Sklar, J. and B. Berkov. 1974. “Abortion, Illegitimacy and the American Birth Rate.” *Science* 13: 909-915.

Sugland, B.W. 1992. “Disparity between Educational Aspirations and Educational Expectations and the Effect on Early Childbearing: Executive Summary.” Unpublished Dissertation, John Hopkins University, Baltimore, MD.

Surveillance, Epidemiology, and End Results (SEER) Program *Populations for 1969-2002.* 2005. National Cancer Institute, DCCPS, Surveillance Research Program, Cancer Statistics Branch.Available at http://seer.cancer.gov/popdata/download.html

Tietze, C. 1973. “Two Years Experience with a Liberal Abortion Law: Its Impact on Fertility Trends in New York City.” *Family Planning Perspectives* 5: 36-41.

Ventura, S.J., T.J. Mathews and B.E. Hamilton. 2001. “Births to Teenagers in the United States, 1940-2000.” National Vital Statistics Reports 49(10), National Center for Health Statistics, Hyattsville, MD.

Wu, L.W. 1996. “Effects of Family Instability, Income, and Income Instability on the Risk of a Premarital Birth.” *American Sociological Review* 61: 386-406.

Wu, L.W. and B.B. Martinson. 1993. “Family Structure and the Risk of a Premarital Birth.” *American Sociological Review* 58: 210-232.

Zabin, L.S. 1994. “Addressing Adolescent Sexual Behavior and Childbearing: Self Esteem or Social Change?” *Women’s Health Issues* 4(2): 92-97.

Table 1
Cross-State *DD* Estimation Results for the Total, African-American and White Teenage Out-of-Wedlock Birth Rates

	TOTAL	
$REPEAL_i \times C70_{jt}$	-	-.0019
		(.0129)
$REPEAL_i \times C7173_{jt}$	-.0115	-.0125
	(.0246)	(.0269)
$REPEAL_i \times C74_{jt}$	-	.0461
		(.0421)
$REPEAL_i \times C7576_{jt}$	-	.0503
		(.0392)
$REPEAL_i \times C7779_{jt}$	-	-.0034
		(.0234)
	AFRICAN-AMERICANS	
$REPEAL_i \times C70_{jt}$	-	-.0306
		(.0371)
$REPEAL_i \times C7173_{jt}$	**-.0967**	**-.1090**
	(.0160)* **	**(.0220)*** **
$REPEAL_i \times C74_{jt}$	-	**-.1049**
		(.0204)* **
$REPEAL_i \times C7576_{jt}$	-	**-.1176**
		(.0218)* **
$REPEAL_i \times C7779_{jt}$	-	**-.1338**
		(.0383)* **

Table 1 (cont.)
Cross-State *DD* Estimation Results for the Total, African-American and White Teenage Out-of-Wedlock Birth Rates

	WHITE	
$REPEAL_i \times C70_{jt}$	-	.0092
		(.0083)
$REPEAL_i \times C7173_{jt}$	.0286	.0310
	(.0271)	(.0269)
$REPEAL_i \times C74_{jt}$	-	**.1087**
		(.0445)* **
$REPEAL_i \times C7576_{jt}$	-	**.0995**
		(.0425)* **
$REPEAL_i \times C7779_{jt}$	-	-.0048
		(.0308)
	t=1982,…,1992	t=1982,…,1998

Notes: Standard errors in parentheses; *** significant at the 1% level. All regressions are clustered at the state level.

Table 2

DDD Estimation Results for the Total, African-American, and White Teenage Out-of-Wedlock Birth Rates

	Total	***African-American***	***White***
1970 Legalization			
$AGEAFFECTED_j \times REPEAL_i \times POST70_t$	**-.0461**	**-.0531**	-.0355
	(.0244)*	**(.0149)*****	(.0367)
(t= 1984,…,1988)			
1973 Legalization (Roe v. Wade)			
$AGEAFFECTED_j \times NONREPEAL_i \times POST73_t$	**-.0993**	**-.1337**	**-.0905**
	(.0100)***	**(.0384)*****	**(.0168)*****
(t= 1987,…,1991)			

Notes: Standard errors in parentheses;* significant at the 10% level; *** significant at the 1% level. All regressions are clustered at the state level.

Figure 1-Cohort Out-of-Wedlock Birth Rates for Total Population between Ages 15-19, 1967-1982

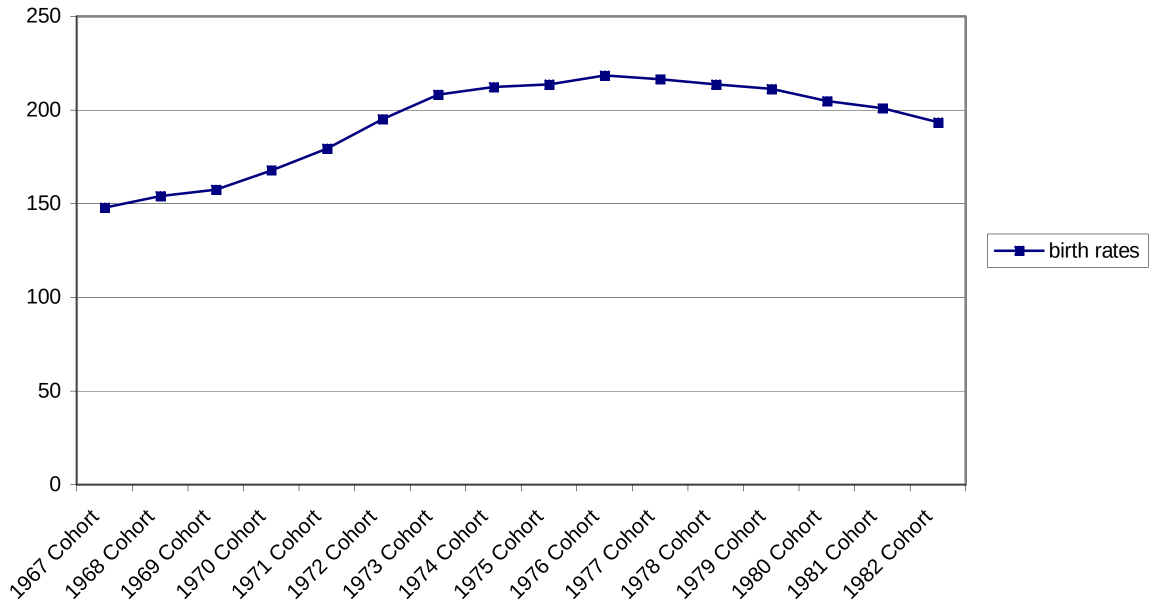


Figure 2- Cohort Out-of-Wedlock Birth Rates for African-Americans and Whites between Ages 15-19,1967-1982

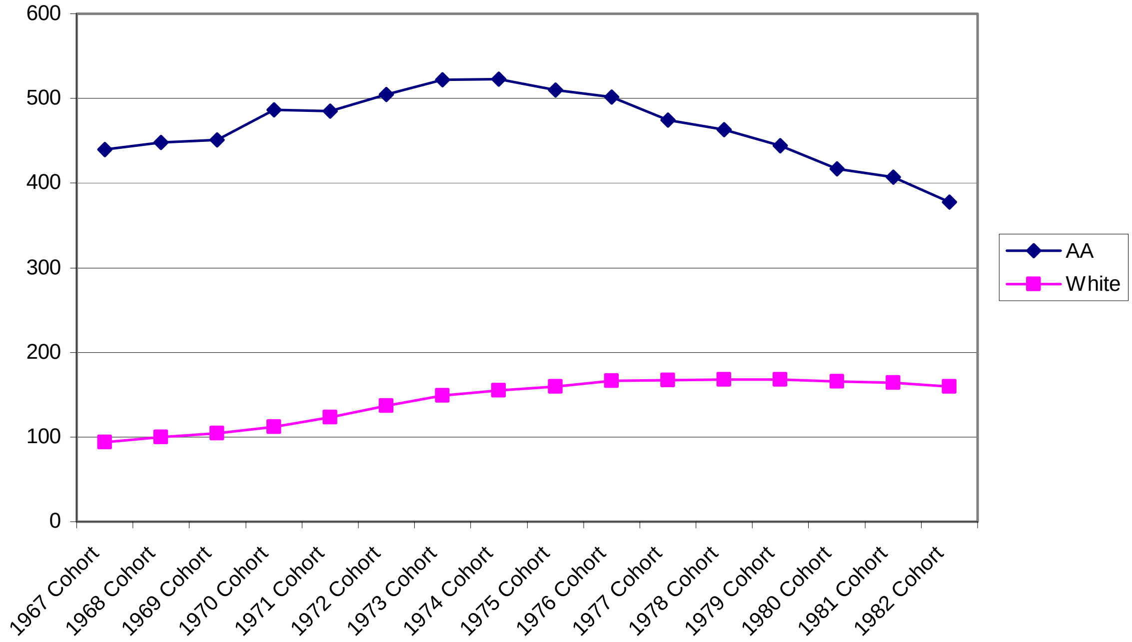


Figure 3 – The Long-Term Impact of the 1970 Legalizations in the Repeal States

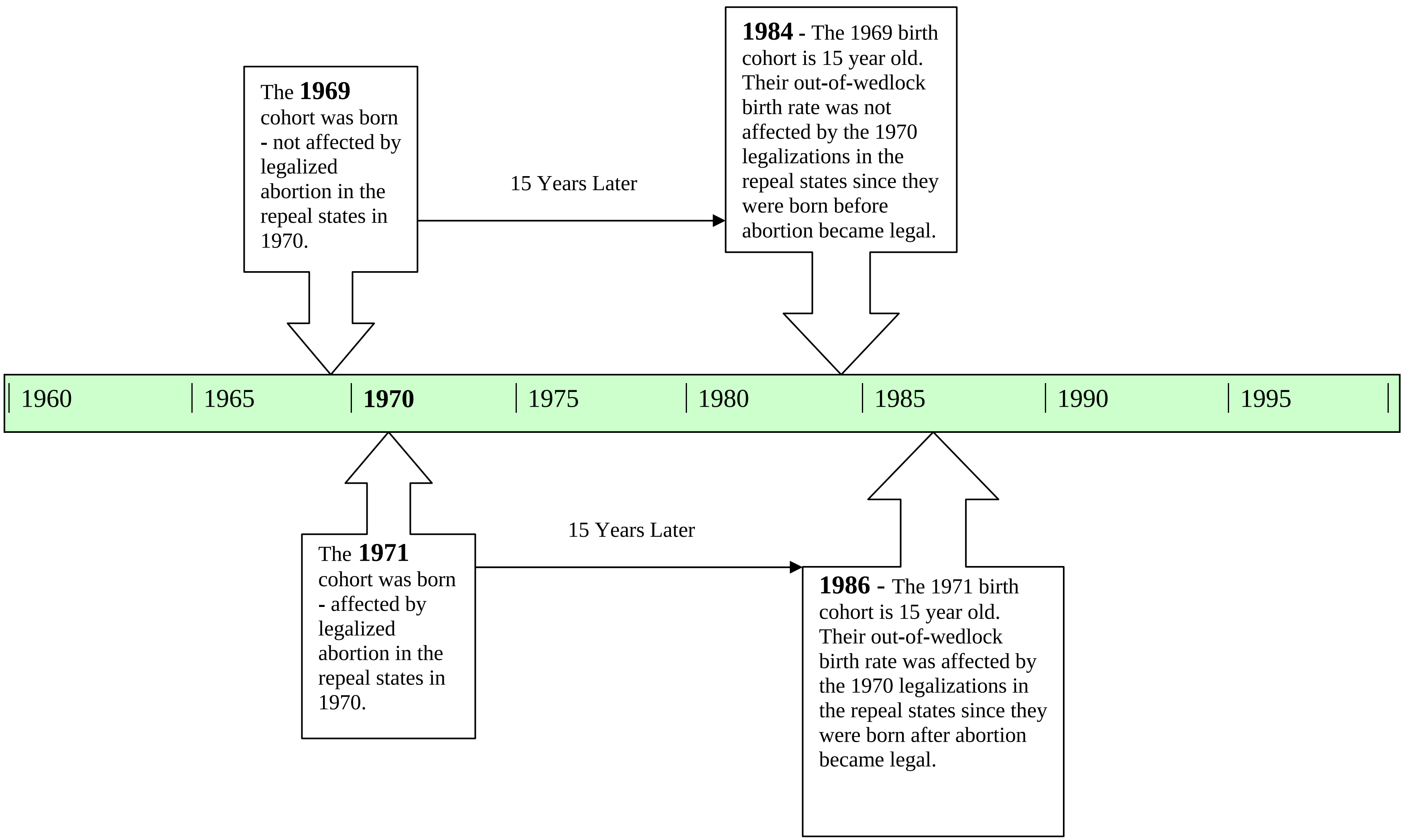
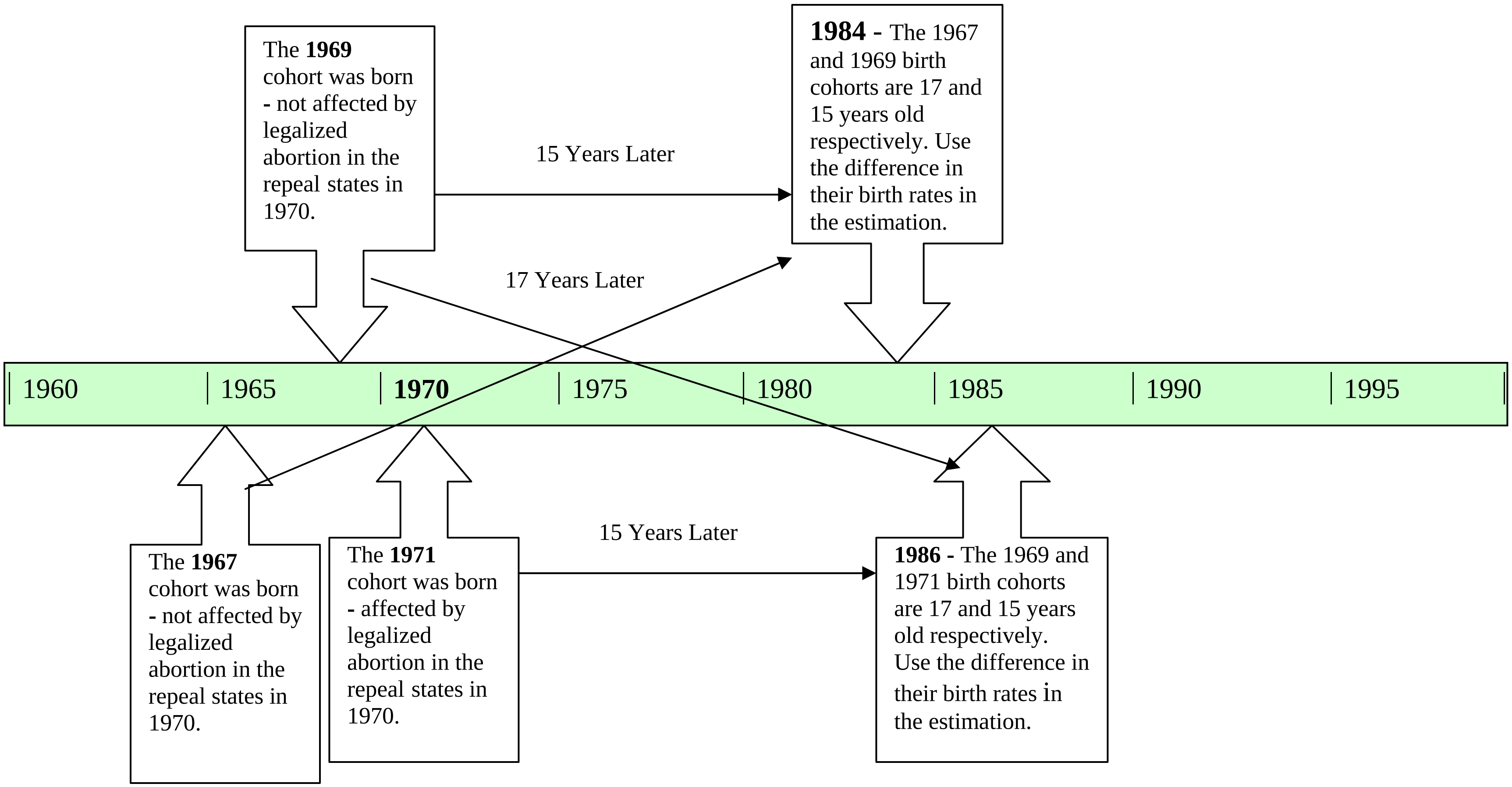


Figure 4 - The Long-Term Impact of the 1970 Legalizations in the Repeal States Using Within-State Comparison Group



Appendix 1

Trends in Teenage Childbearing

In this appendix we present the basic trends for teenage childbearing in the United States. In Figure A1, we present the trend for adolescent childbearing in general, without making any distinction between births to unmarried or married mothers. We then consider the trend for teenage out-of-wedlock childbearing in Figure A2. Furthermore, because the general trend for all teenagers hides important differences between Whites and African-Americans in levels as well as in trends, we offer the respective figures for these two groups in Figure A3 and A4.

The rate of teenage childbearing in the United States has declined precipitously since the late 1950's, except for a brief, but steep, rise in the late 1980's through 1994. The birth rate peaked in 1957 at a rate of 96 births per 1000 women aged 15-19, and then declined to its lowest level of 37 births in 2001 (Ventura et al., 2001)[38]. This general trend is a reflection of the birth rate for Whites, who comprise majority of the population, but the corresponding trend for African-Americans is similar, and in fact the gap between the birth rates for the two races shrank over this period. In 1960, the African-American birth rate was 156 while it was 79 for Whites. By the year 2001, the birth rate declined to 72 for African-Americans and 41 for Whites.

While teenage childbearing has been in a long-term decline in the United States since the late 1950s, the story for the non-marital birthrate among teens is sharply different. In 1957 when total teenage child bearing was its highest all-time high level, the

[38] Throughout this paper we will be using the birth rate per 1000 women in a specified group as our childbearing measure for teenagers. Consider, for example, African-American teenagers. The birth rate for this group will be calculated as the ratio of the total number of births to African-American women aged 15-19 to the total number of women in this age group for given year (and state when we need state specific information).

birth rate per 1000 unmarried women aged 15-19 was 16. Almost four decades later, the out-of-wedlock birth rate hit its maximum of 46 births in 1994. Between 1994 and 2001 it declined to 37 births. For African-Americans, a steep increase occured in the out-of-wedlock birth rate from 1960 to 1991, followed by a period of constant decline. For Whites, there was a persistent increase (although to a lesser degree than African-Americans) from 1960 to 1994, and then a decline for the following decade (Figure A4). Even though the increase of unmarried teen mothers has drawn considerable attention in academic circles, the reasons behind this reversal in the teen out-of-wedlock birth rate in the early 1990s have not been investigated by many studies and are still ambiguous. Darroch and Singh (1999) analyzed the reasons behind the recent decline in the U.S. teen pregnancy rate, using data from two comparable, large-scale government surveys, the 1988 and 1995 cycles of the National Surveys of Family Growth, as well as recent information on the rates of teenage pregnancy, births and abortions. Their analysis concluded that approximately one-quarter of the decline in teenage pregnancy in the United States between 1988 and 1995 was due to increased abstinence. (The proportion of all teenagers who had ever had sex decreased slightly, but non-significantly, during this period, from 53 percent to 51 percent) Approximately three-quarters of the decline resulted from changes in the behavior of sexually experienced teens. (The pregnancy rate among this group had fallen percent, from 211 per 1,000 to 197.)

Darroch and Singh (1999) considered a number of behavioral changes that could explain why a smaller proportion of sexually experienced teenage women became pregnant in 1995 than in 1988, including the possibility that they were having less sex. However, they found that there was little change overall between the two years in how often sexually experienced teenagers had intercourse. Instead, the researchers found that overall contraceptive use increased—but only slightly, from 78 percent in 1988 to 80 percent in 1995. More importantly, teenagers in 1995 were choosing more effective methods (Boonstra, 2002).

Figure A1-Total Birth Rate for Women Aged 15-19 in the U.S., 1960-2001

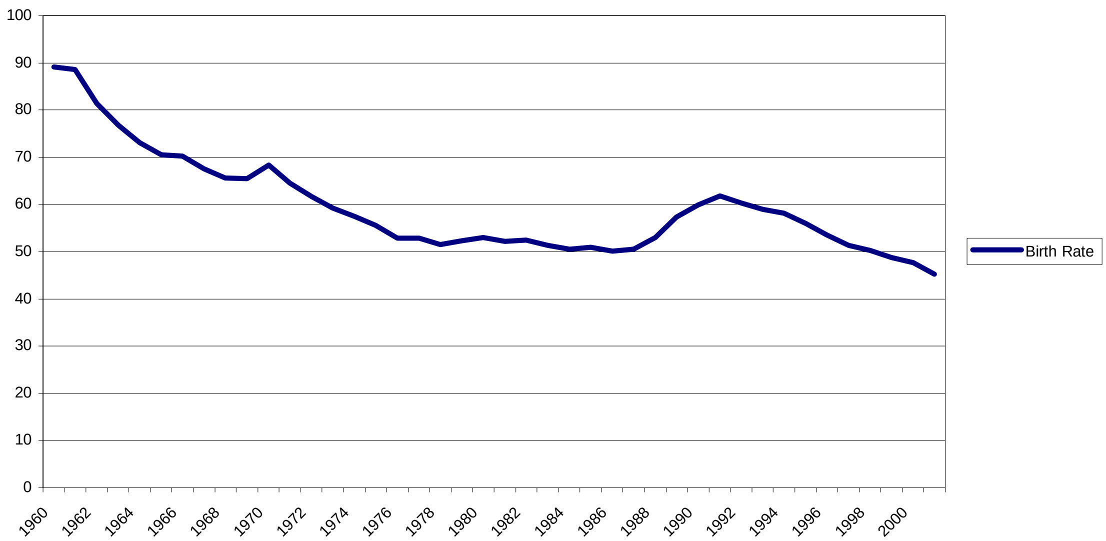


Figure A2 - Out-of-Wedlock Birth Rates for Women Aged 15-19 in the U.S., 1960-2001

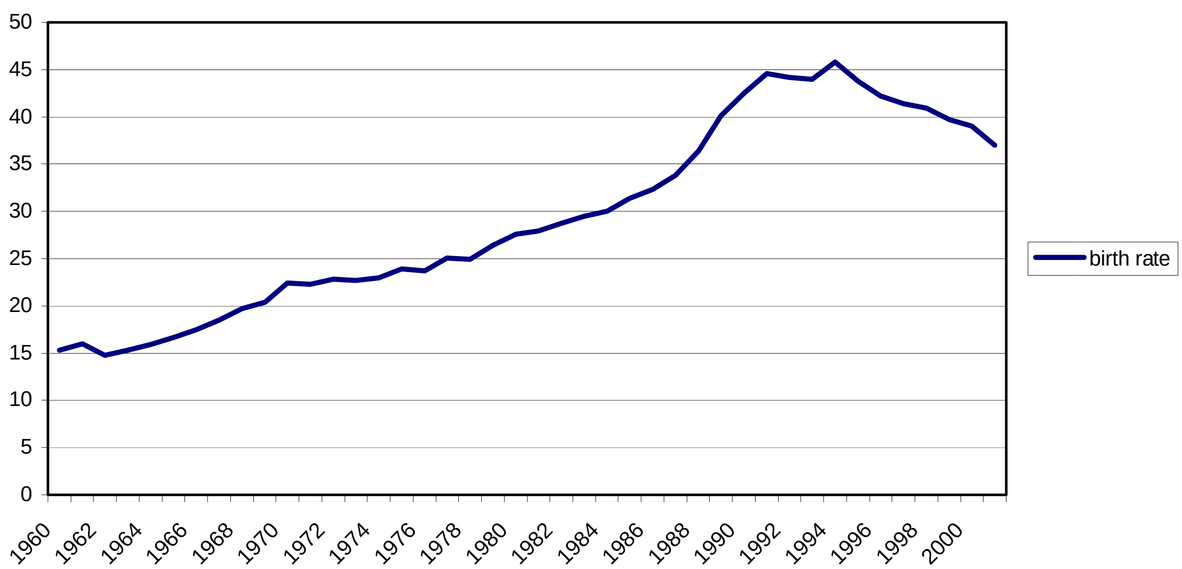


Figure A3- Birth Rates for African-American and White Women Aged 15-19 in the U.S., 1960-2001

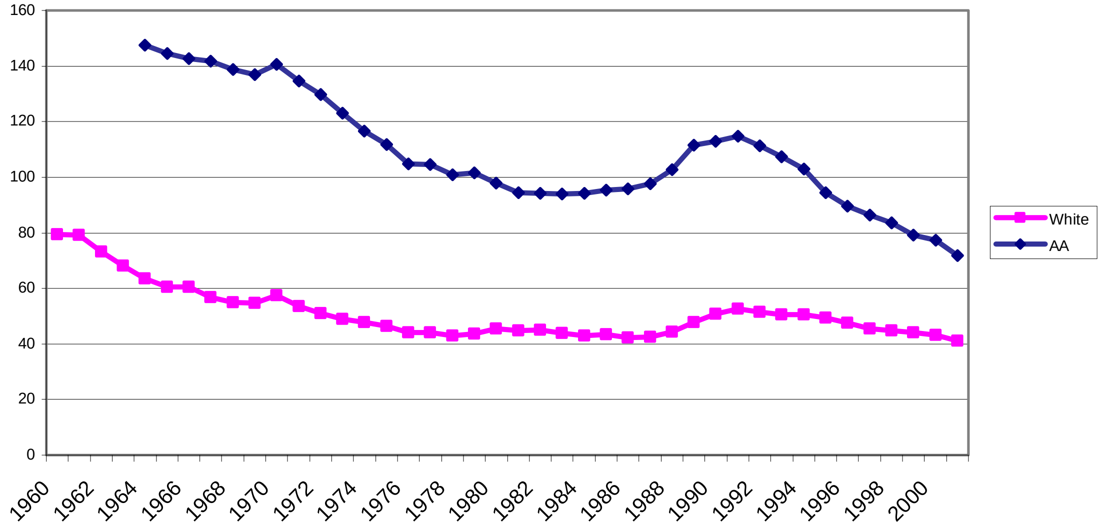
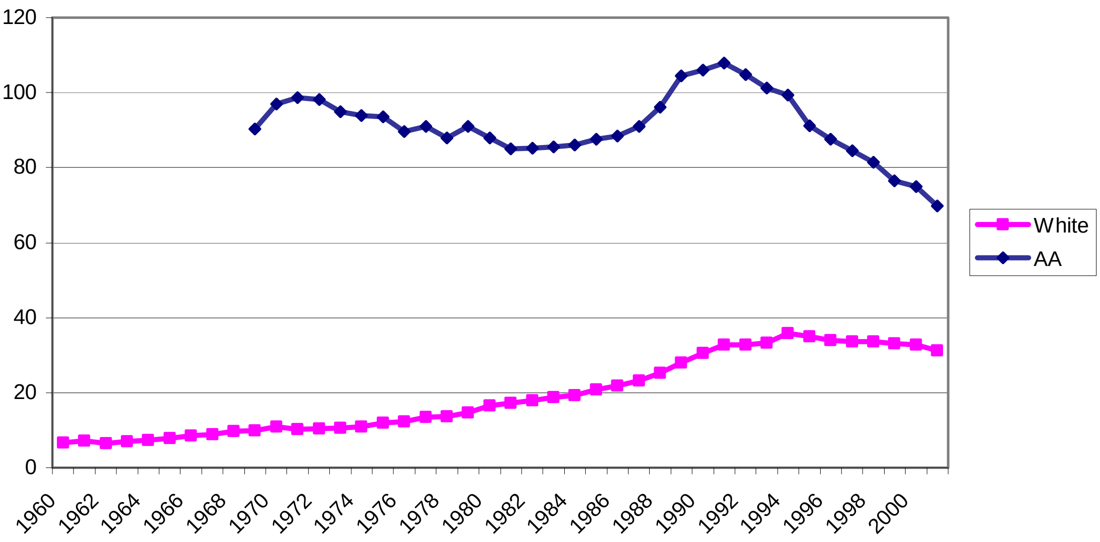


Figure A4 - Out-of-Wedlock Birth Rates for African-American and White Women Aged 15-19 in the U.S., 1960-2001



Appendix 2

In this appendix we show that our coefficient of interest is indeed the corresponding *DDD* parameter. As we discussed in the methodology section, *DDD* parameter-suppose for simplicity we are only interested in the impact on 15 year-olds-can be written in the following way

$$\begin{aligned}
\delta_5 &= \underbrace{(\Delta\overline{BR}_{(r)jt} - \Delta\overline{BR}_{(nr)jt})}_{DD_{affected}} - \underbrace{((\Delta\overline{BR}_{(r)j't} - \Delta\overline{BR}_{(nr)j't}))}_{DD_{unaffected}} \\
&= \underbrace{((\overline{BR}_{(r),15,1986} - \overline{BR}_{(r),15,1984}) - (\overline{BR}_{(nr),15,1986} - \overline{BR}_{(nr),15,1984}))}_{DD_{affected}} \qquad (A1) \\
&\quad - \underbrace{((\overline{BR}_{(r),17,1986} - \overline{BR}_{(r),17;1984}) - (\overline{BR}_{(nr),17,1986} - \overline{BR}_{(nr),17,1984}))}_{DD_{unaffected}}
\end{aligned}$$

where r and *nr* subscripts indicate the repeal and nonrepeal states respectively. Note that we use 17 year-olds as the unaffected comparison group. Also, we can write our regression without logarithmic transformation of the birth rates[39] as

$$\begin{aligned}
\ln BR_{ijt} &= \delta_0 + \delta_1 AGEAFFECTED_j + \delta_2(AGEAFFECTED_j \times POST70_t) \\
&\quad + \delta_3(AGEAFFECTED_j \times REPEAL_i) + \delta_4(REPEAL_i \times POST70_t) \qquad (A2) \\
&\quad + \delta_5(AGEAFFECTED_j \times REPEAL_i \times POST70_t) + s_i + y_t + e_{ijt}\,.
\end{aligned}$$

Let us write each parameter in *DDD* expression in terms of the coefficients that capture particular birth rates

$$\overline{BR}_{(r),15,1986} = \delta_0 + \delta_1 + \delta_2 + \delta_3 + \delta_4 + \delta_5$$

$$\overline{BR}_{(r),15,1984} = \delta_0 + \delta_1 + \delta_3$$

$$\overline{BR}_{(r),17,1986} = \delta_0 + \delta_4$$

[39] Clearly, the coefficients will be different without the logarithmic transformation but we keep the same notation for the coefficients in this appendix for simplicity.

$$\overline{BR}_{(r),17,1984} = \delta_0 \tag{A3}$$

$$\overline{BR}_{(nr),15,1986} = \delta_0 + \delta_1 + \delta_2$$

$$\overline{BR}_{(nr),15,1984} = \delta_0 + \delta_1$$

$$\overline{BR}_{(nr),17,1986} = \delta_0$$

$$\overline{BR}_{(nr),17,1984} = \delta_0$$

Now rewriting and rearranging the *DDD* equation given this information will yield

$$\begin{aligned} DDD &= ((\delta_0 + \delta_1 + \delta_2 + \delta_3 + \delta_4 + \delta_5 - \delta_0 + \delta_1 + \delta_3) - (\delta_0 + \delta_4 - \delta_0)) \\ &\quad -((\delta_0 + \delta_1 + \delta_2 - \delta_0 - \delta_1) - (\delta_0 - \delta_0)) \\ &= (\delta_2 + \delta_4 + \delta_5 - \delta_4) - \delta_2 \\ &= \delta_5, \end{aligned} \tag{A4}$$

which is our coefficient of interest.

Appendix 3

Figure A5- Abortion Rates per 1000 Live Births in the Repeal and Non-Repeal States, 1970-1985