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**Did Welfare Reform Differentially Influence the Fertility of Young Teens?
Assessing the Success of the “Living Arrangement Rule”
and the “Stay-in-School Rule”**

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Abstract

The Aid to Families with Dependent Children program was designed to provide cash assistance primarily to fatherless children regardless of the age of the child’s mother and her living arrangements. During the 1990s several states received and implemented waivers to their programs requiring teenage mothers younger than 18 to live with an adult parent or legal guardian and enroll in high school in order to receive benefits. Later Congress added these rules to the Personal Responsibility and Work Opportunity Act of 1996 making them applicable to teen mothers who enrolled in the Temporary Assistance for Needy Families Program. We examine whether the “living arrangement rule” and “stay-in-school rule” reduced fertility among women under 18 using natality data from the National Center for Health Statistics. Our findings suggest that these rules led to a decline in fertility, primarily affecting non-Hispanic African American and non-Hispanic white teenagers. Our estimates imply that the percent decline in fertility rates post-reform was 9.2 percentage points larger for teens aged 15 to 17 than for teens aged 18 or 19.

Key words: teenage childbearing, welfare reform

With the passage of the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 (PRWORA), one of the primary components of the social safety net in the United States changed dramatically. PRWORA eliminated the Aid to Families with Dependent Children (AFDC) program and created a new block grant for low-income families called Temporary Assistance for Needy Families (TANF). Unlike the AFDC program, TANF is not an entitlement and states were given considerable discretion in determining the form of assistance and eligibility. The TANF program also includes strict work requirements for recipients with harsh sanctions for those who do not comply

and a lifetime time limit of 60 months for program receipt (states do have the authority to exempt some recipients).¹

In addition to the changes in funding and the new emphasis on self-sufficiency through work, Congress made another striking change in the means-tested cash transfer system with TANF. In 1935 when the AFDC program (initially called Aid to Dependent Children) was created, the program's primary purpose was to provide cash assistance to needy, fatherless children. In contrast, PRWORA made explicit the objective to use TANF to influence the family life of low-income families. This emphasis on low-income families is evident in the goals stated in PRWORA:

- (1) provide assistance to needy families so that children may be cared for in their own homes or in the homes of relatives;
- (2) end the dependence of needy parents on government benefits by promoting job preparation, work, and marriage;
- (3) prevent and reduce the incidence of out-of-wedlock pregnancies and establish annual numerical goals for preventing and reducing the incidence of these pregnancies; and
- (4) encourage the formation and maintenance of two-parent families (U.S. Public Law 193, Section 401).

Within this major focus on family structure, unwed teenagers less than 18 years old received particular attention. In the findings section of PRWORA, Congress notes the high proportion of unwed mothers younger than 18 who receive public assistance.

¹ Information in the opening paragraph comes from Moffitt (2003). See this source for a detailed description of the changes made with the TANF program.

Congress also lists the negative outcomes associated with out-of-wedlock childbearing.² To receive benefits, these young women were now required to attend school or a training program unless they had completed high school (hereafter referred to as the “stay-in-school rule”).³ TANF also forbids providing aid to these mothers unless they live “...in a place of residence maintained by a parent, legal guardian, or other adult relative of the individual as such parent’s, guardian’s, or adult relative’s own home” (U.S. Public Law 193, Section 408; hereafter referred to as the “living arrangement rule”). These minor parents rules were actually first implemented in 1993 in Michigan after the state received a waiver for the AFDC program. Shortly thereafter, several other states followed suit and by the time that PRWORA was passed, sixteen states had implemented these minor parent rules. Since these welfare reforms reduced the net benefit of a nonmarital birth compared to the benefits available in the AFDC program, one might expect welfare reform to reduce nonmarital fertility, all else equal (Horvath-Rose and Peters 1999, 2001; Kaestner, Korenman, and O’Neill 2003; Peters, Plotnick, and Jeong 2001).⁴

Given the major emphasis Congress placed on reducing childbearing among minors with PRWORA, surprisingly little research has been completed on the topic. This paper attempts to address the paucity of research in the area. Our objective is to determine if the changes created by welfare reform reduced the childbearing among women under age 18. More specifically, we estimate if these changes had a differential

² By 1996, the vast majority of births to teenagers was nonmarital (75.9 percent) (Ventura and Bachrach 2000). Among African Americans, nearly all teen mothers were unmarried (95.4 percent), while 69.3 percent of non-Hispanic white and 67.7 percent of Hispanic teen mothers were unmarried (Ventura and Bachrach 2000).

³ Congress added the additional stipulation that the stay-in-school rule only applied once the child of the teenager reached 12 weeks of age.

⁴ It is noteworthy that 44 states do allow teen mothers to live alone with state approval (Boonstra 2000). We have found no information describing the difficulty in obtaining state approval or if teens are aware of this stipulation.

impact for “young teens,” which we define as those aged 15 to 17, compared to “old teens,” or those 18 and 19.

Understanding the impact of welfare reform is crucial. From a research perspective, we hope to improve our understanding of what Rebecca Blank (1997) calls the “independence effect” of welfare: there is some evidence that young women valued the freedom to move out of their parents’ home provided by the resources available through AFDC. If true, we should find that removing this benefit reduced teenage fertility. From a policy perspective, this research strives to document if the removal of this benefit altered teenagers’ fertility decisions post-PRWORA thereby contributing to the debate on the success of the TANF program.

As Congress argued in the preamble to the PRWORA, reducing teenage childbearing might improve a variety of deleterious outcomes for both teen mothers and their children. We know that having a child at a young age is associated with a reduction in educational attainment for teen mothers, particularly among those younger than 18, as well as with poverty, welfare receipt, and a variety of risky behaviors (Geronimus and Korenman 1992; Hoffman, Foster, and Furstenberg 1993; Klepinger, Lundberg, and Plotnick 1995; Levine, Pollack, and Comfort 2001).⁵ If these additional requirements for young teenage mothers reduce the likelihood that young women will bear children, we might see an improvement in a variety of outcomes.

⁵ It is noteworthy that there is no consensus regarding the consequences of teenage childbearing. In early work using only cross-sectional data, researchers typically noted the strong correlation between teenage childbearing and negative outcomes for both teenage parents and their children (Hoffman 1998; Levine, Pollack, and Comfort 2001). Recently, researchers have begun to address the selection of mothers into teenage childbearing, asking if it is teenage childbearing per se or the unobserved heterogeneity among the mothers who have teen births that leads to deleterious consequences. The latest evidence suggests that teenage childbearing is not as detrimental as first thought and that studies based on cross-sectional data lead to upwardly biased estimates of the effect of teenage childbearing. Thus, the jury is still out regarding the causal impact of teenage childbearing.

We use data from the National Center for Health Statistics (NCHS) natality series, population estimates from the U.S. Bureau of the Census, as well as a variety of other sources to identify the role welfare reform played, if any, in reducing teen fertility during the 1990s. Our results suggest that compared to older teenagers, the percent change in fertility rates among young teenagers was about 9.2 percentage points lower in the post-reform period. Further, these differences were largest among African Americans and white teenagers.

This paper proceeds as follows: in the next section we describe the theoretical reasons why TANF may have had a differential impact on younger teenagers and review the relevant literature. Next, we describe the data we use in this analysis followed by a description of our analytical plan. In the subsequent section, we report results before concluding in the final section.

2. Background and Literature

Economists have long made theoretical arguments that providing welfare benefits only to low-income single mothers creates incentives for women to delay marriage, divorce or separate, postpone remarriage, and bear children out-of-wedlock (Moffitt 1992). Since the welfare reforms in the 1990s unambiguously reduced the net benefit of the program with, among other things, the introduction of time limits, work requirements, and the removal of the entitlement, coupled with declining real benefits in almost all states (U.S. Congress 2004), one should expect, all else equal, that the proportion of low-income women who had a nonmarital birth declined following welfare reform.

Economic theory suggests that the impact of welfare reform is unambiguous for teenagers as well. Imagine a world in which a group of teenagers was offered a package of welfare benefits consistent with those offered in the U.S. in 1990 and each was weighing the decision to become pregnant and have a child. Given the costs and benefits the teens perceived, some portion of the teenagers would decide to have the child. Now imagine that the exact same group of teens was making the decision in the post-TANF regime. These potential-mothers would face a five-year limit on benefits and, in some instances, a family cap might reduce the incentive to have additional children while receiving benefits (Horvath-Rose and Peters 2001; contra Kearney 2004). In this regime in order to receive cash transfers these young women would have to complete high school. If they were younger than 17 and unmarried, they could not live separate from their parents, guardians or other adult relatives as well.

The teenagers who chose not to have a child in the first scenario would not alter their decision in the second situation since the benefits of bearing a child were even lower. Some proportion of the teenagers who chose to have a birth under the first scenario, i.e., those near the margin, would no longer find that the benefits of having a child outweighed the costs and would decide not to bear a child. Given this argument, therefore, we expect a decline in teenage births for all young women aged 15 to 19 post-PRWORA with an even larger decline for teenagers under the age of 17. We know that about three-quarters of all unwed teenage mothers received AFDC within five years of the birth of their child (Adams and Williams 1990, cited in Wertheimer and Moore 1998). Therefore, welfare reform may have had a large impact on teenage childbearing in the United States.

Although it has not received a considerable amount of research attention, there is some evidence that the living arrangement rule and the stay-in-school rule are being enforced. In non-random surveys completed in Atlanta and Chicago, Shapiro and Marcy (2002) report that approximately 40 percent of teen mothers interviewed who were not receiving TANF had applied and been turned down. Over half of the denials were classified as failing to comply with school requirements, not living with an adult, or “other” (which does not include failing to meet the income requirements). Further, among those who received benefits, about one-fifth had their benefits reduced (or had lost them entirely) mainly due to missing school or an appointment with a case worker.

Duffy and Levin-Epstein (2002) use data from a 2000 survey from the Center for Law and Social Policy (CLASP) asking states about their enforcement of the living arrangement rule and the stay-in-school rule. Eleven states responded to their inquiry about TANF and teenagers. It is hard to generalize to the entire population of teenagers given the limited response, but Duffy and Levin-Epstein claim that the states that responded accounted for 49 percent of all teen births in March of 1999.

Only four states responded to a question asking if teen mothers who were in violation of the stay-in-school rule were denied access to the program. All four responded no. Two states responded to questions asking if violations of the stay-in-school rule were the cause for closing a case. Arizona reported that 19 percent of those cases subject to this rule were closed due to a violation, while Illinois reported hardly any were closed for this reason. According to this survey, however, teens who were in violation of the stay-in-school rule were frequently sanctioned. Of all teens who were subject to the stay-in-school rule, 5.6 percent of those in California were sanctioned, 6.4

percent of those in Arizona, 9.9 percent of those in Ohio, 10.5 percent of those in Illinois, and 22.7 percent of those in Texas were sanctioned (only five of the 11 states responded to this question). Duffy and Levin-Epstein argue that these proportions were much higher than the overall sanction rate for TANF recipients in the state. They also report that between one-third and one-half of all sanctions were for second or higher violations.

Duffy and Levin-Epstein report that few states reported the number of teenagers subject to the living arrangement rule in their survey. They cite work from another CLASP study which shows that among the nine (of the 50) states that responded to this other survey, between 0 and 7 percent reported that a teen was declared ineligible for TANF due to a violation of the living arrangement rule. They do not provide any information on sanctions for violations of the living arrangement rule.

In another study, Kalil and Danziger (2000) found that most (approximately 60%) teenage mothers understood that welfare benefits depended on their complying with the living arrangements rule and/or the stay-in-school rule even before they gave birth (regardless of whether they actually enrolled in the welfare program). Of course, Kalil and Danziger's findings may not be generalizable to the entire population of teenagers either: they use a sample of teenage mothers from one county in Michigan.

The literature on the impact of welfare on family structure is extensive and ambiguous. Economists and demographers have published a variety of studies showing different effects for AFDC on teenage fertility depending on the data, the empirical methods, and racial/ethnic group being studied (Kearney 2004). Several studies show no statistically significant effect of AFDC (e.g., An, Haveman, and Wolfe 1993; Mayer 1997; Duncan and Hoffman 1990; Hoffman and Foster 2000; Hoynes 1997). Others

show positive effects, such as Plotnick (1990) who finds weak evidence of a positive link between AFDC and fertility for white and African American teenagers but no effect for Hispanics and Lundberg and Plotnick (1995) who find a strong positive association for white teens but no effect for African American or Hispanic teens.

More recently, however, the analyses have become more sophisticated and are more consistently finding welfare effects for marriage and fertility (although this literature is not confined to teenagers). Summarizing this broad literature, Moffitt (2003) writes,

A simple unweighted average of the estimates, therefore, reveals a central tendency suggesting the presence of an effect of welfare benefits on family structure. Moreover, when distinctions are made between the studies by the likely credibility of estimates – those controlling for the most other variables, and which concentrate on the most appropriate part of the distribution of women, and use the more careful econometric methods – effects are sometimes stronger, sometimes substantially so.... Therefore, although there is still considerable uncertainty in the literature and there remain a large number of studies reporting insignificant estimates, this reading of the literature leads to the conclusion that welfare is likely to have some effect on family structure (p. 336).

There are a handful of studies investigating the relationship between welfare reform and/or TANF and teenage fertility. Since these are the studies more relevant to this work, below, we describe these studies in detail. We also explain how this paper builds on this previous literature.

Horvath-Rose and Peters (1999, 2001) investigated the relationship between a variety of welfare waivers and the teenage nonmarital birth ratio (the number of teen births to single mothers/number of teen births) using data from the NCHS natality series from 1984 to 1996. Their results showed that states with waivers requiring teenagers to live at home and attend school had higher nonmarital birth ratios, all else equal, compared to states without the waivers: a result that was different from their expectations a priori. They explain that the waiver may have offered a measure of security for the teen mother. If her home situation was adequate, the teen could reside there. If not, the teen could move into a supervised group home.

Horvath-Rose and Peters did not separate 15- to 17-year-olds, the group that was targeted by this policy, from the 18-19 year-olds, and, of course, their outcome measure is different from the outcome used in this study. A nonmarital birth ratio can increase because the number of births falls holding the number of nonmarital births constant, because the number of nonmarital births increase holding the number of births constant, or because both the numerator and denominator changed. We concentrate only on the number of teenage births, their denominator. Another difference is the time period under investigation. Only a handful of states implemented these rules for minors prior to 1995, while all states had implemented TANF by 1997. This study, therefore, builds on their work by adding data post-PRWORA: a period with more variation in the policy both within and between states and separating the analysis by the age of the teenagers, all of which may lead to different results.

Using data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY) and the 1997 cohort of the NLSY, Kaestner et al. (2003) find evidence that

welfare reform reduced teenage fertility for a sample of 17-year-olds and a sample of 19-year-olds. More specifically, for both age groups they estimate the difference in the probability of a teen birth for those in the 1997 cohort, a group they argue had been impacted by the 1996 welfare reform, compared to teens in the 1979 cohort, a group making fertility decisions, obviously, well before the reforms of the 1990s. Within each age group they also separated their analysis into an advantaged group and a disadvantaged group based on the education of the teen's mother and the family structure of the daughter when she was either age 14 (1979 cohort) or 12 (1997 cohort). They find that welfare reform reduced the probability of a teen birth for both age groups in the "high risk group," although it was only statistically significant for the 19-year-olds.

Kaestner et al. (2003) correctly acknowledge that unobserved factors may be biasing the estimates of welfare reform on teenage fertility. To address the potential for omitted variable bias, Kaestner et al. use a difference-in-differences (DD) model subtracting the estimated coefficient for reform for the high risk group (post-TANF – pre-TANF) from the estimated coefficient for reform for the low-risk group. They argue that if these omitted factors bias the coefficients for both groups equally, then taking the difference would produce an unbiased estimate of the impact of welfare reform on teenage fertility for the high-risk group. Their DD results show large and statistically significant reductions in the probability of a teen birth due to welfare reform for teens of both ages. Although their results suggest that welfare reform did reduce teenage fertility, it is difficult to determine if females under the age of 17 were more affected than teenagers 18 and 19 from their data since they use cumulative probabilities for both the

17- and 19-year-olds. In other words, it is impossible to ascertain whether a 19-year-old mother had her child before or after the age of 18.

This paper adds to the work of Kaestner et al. (2003) in a number of ways. First, we are more interested in determining if there was a differential decline in teenage childbearing among young teenagers rather than all teenagers. Therefore, we are not using cumulative probabilities as was the case with Kaestner et al. As a result, we can compare birth rates for teens under age 18 to teens aged 18 and over to estimate if welfare reform had a differential effect for young teens. We also have information on the fertility of teens aged 15 through 19, whereas Kaestner et al.'s results are confined to 17 and 19-year-olds. Teenagers, particularly young teenagers, often do not consider the long-term impact of their behaviors (Leibowitz, Eisen, and Chow 1986). Failure to include this group when considering policy changes may tend to overstate the impact of welfare reforms. Finally, by using the National Center for Health Statistics Natality Data, we have nearly all recorded births to teenagers in the United States every year making biases due to attrition, undercoverage, inaccurate response, or nonresponse less likely (See Moffitt and Ver Ploeg (2001) for a discussion of potential problems with using the NLSY surveys for welfare inquiries).

Hao and Cherlin (2004) also use data from the NLSY97. Using an event history model, they follow two cohorts of adolescents from age 14 to 16: one pre-welfare reform (born in 1980) and one post-welfare reform (born in 1982 or 1983). Similar to Kaestner et al. (2003), Hao and Cherlin use a DD model comparing estimates for females in a high risk group (defined in a variety of ways including poverty status, AFDC receipt, and low

parental education) and a low risk group before and after welfare reform. Their estimates suggest a weak, positive relationship between welfare reform and teenage childbearing.

Hao and Cherlin do not include waiver data in their analysis. If the teens who resided in states with welfare waivers responded to the reforms, their “control group” would be contaminated and any differences between the treatment and control would be minimized. Second, due to data limitations, Hao and Cherlin censored their data once the young women reached age 16. Given the young age at which their event history was censored, Hao and Cherlin’s results may not be applicable to the vast majority of teens. In 1999, there were over 480,000 births to young women between the ages of 14 and 19. Of those births, less than 17 percent occurred among those aged 14 to 16 (authors’ calculations using Natality data from 1999). Hao and Cherlin acknowledge this potential problem. They write, “Although our design has allowed adolescents time to learn the new rules and norms from age 14, these new rules and norms may not come to their awareness until later ages. Therefore, our analysis may have missed the actual effect of welfare reform” (p. 192). Our panel includes births to all teenagers aged 15 to 19, the ages that include the vast majority of all teenage births. Finally, the criticisms levied by Moffitt and Ver Pleog (2001) regarding the NLSY97 apply here as well.

Acs and Koball (2003), again using data from the NLSY97, compared the difference in the likelihood of childbearing for 15- to 17-year-olds in the first interview year (January to September, 1997) to the fertility of 15- to 17-year-olds in the third round of the NLSY97 (interviewed between October, 1999 and April, 2000). They find a decline in birth rates of 20 percent, but the difference was statistically insignificant.

Acs and Kobel do not provide a lot of detail for the model that they used other than to say that the distribution of teens by age was not uniform so their results were “regression adjusted” for age. Thus, it appears that unmeasured factors that may bias the coefficients for teenage childbearing are an issue to consider. They are also using a nationally representative sample so the criticisms levied by Moffitt and Ver Ploeg (2001) are relevant here as well. Finally, the “pre-PRWORA” cohort may have been considering having children after TANF was implemented in many states and thus does not represent a “treatment” group even in a quasi-experimental sense.⁶

A second related and much smaller literature examines the autonomy unmarried teens could gain from their parents through the cash benefits provided by the AFDC program. For young women with a strong preference to establish independence from their parents or for those living in abusive homes, this potential independence, conditional on a nonmarital birth, was an important benefit of the AFDC program. In fact, evidence for the “independence effect” of welfare may be a bit stronger than the evidence for a fertility effect (Blank 1997). For instance, young single mothers were more likely to co-reside with their parents in low AFDC-benefit states, while young single mothers in high benefit states were much more likely to live separate from their parents (Blank 1997; Ellwood and Bane 1985). Ellwood and Bane (1985) estimate that a one standard deviation increased in AFDC benefit levels for a non-White, single mother aged 20 without a high school degree increased the probability that she lived independent from her parents from 22 percent to 43 percent. Among a sample of women 36 or younger, Hutchens, Jakubson, and Schwartz (1989) find that the greater the gap between

⁶ Acs and Kobel argue that since the decisions to have a child occurred six to nine months before they were interviewed, this is less of an issue.

welfare benefits for mothers who live without other adult family members and the benefits provided to a mother who co-resides with other family members, reduces the likelihood that the mother receives welfare and lives with her family relative to the probability that she receives welfare and lives alone. These results were not statistically significant, however. If increases in welfare benefits do increase the probability that young mothers live alone, then we expect the reduction in the ability to gain independence created by welfare reform in the mid 1990s to deter young teens differentially compared to older teens.

Bitler, Gelbach, and Hoynes (2004) use the CPS to examine the effect of welfare waivers and TANF on the living arrangements of children. They find that, among African-American children, TANF implementation is associated with a large (2 to 4 percentage points) reduction in the fraction living with an unmarried parent, an increase in the fraction living with a married parent, and an increase (3 to 7 percentage points) in the fraction of these children living with neither parent.

Data

To measure teen fertility rates, we use a couple of data sources. First, we use data from the NCHS detailed natality series to obtain population data on the number of births to females aged 15, 16, 17, 18, and 19 by race/ethnicity i.e., non-Hispanic white (hereafter referred to as “white”), non-Hispanic African American (hereafter referred to as “African American”), Hispanic (which includes Hispanic white and Hispanic African American), Asian or Pacific Islander (both Hispanic and non-Hispanic), and American Indian (both

Hispanic and non-Hispanic) in each state from 1992 to 1999.⁷ We supplement these data with the corresponding annual state population estimates by age, race/ethnicity, state, and year from the U.S. Bureau of the Census from 1992 to 1999. We generated age*race/ethnicity*state*year-specific teen birth rates by dividing the total number of births to females of a given age and race/ethnicity in each state each year by the total number of women of the same age and race/ethnicity in each state each year (*1000). These data sources are reported by Hamilton, Sutton, and Ventura (2004) in the National Vital Statistics Reports, and with our data, we duplicate the rates they report.⁸ In addition to these data, we cull state level data from a variety of sources.

We use an indicator variable to represent the implementation date for the minor parent rules (both the stay-in-school rule and the living arrangement rule). We set this variable equal to one in the year following the official date of implementation for the sixteen states that implemented these rules after receiving a waiver. For the remaining states we set the indicator to one in the year nine months following the implementation date for TANF. In the rare instance, when the actual date of implementation differed from the official date of implementation, we use the actual date.⁹ In addition we use the welfare benefit for a family of four (inflated to 1999 dollars using the CPI-U-X1), state unemployment rates, and abortion policies as controls. Table 1 contains descriptive statistics for our data.

⁷ Following Hamilton, Sutton, and Ventura (2004), we group Aleuts and Eskimos in the American Indian category.

⁸ One exception is the Hispanic category. Hamilton, Sutton, and Ventura classify Hispanics as any individual regardless of race who reports being Hispanic. Given the small numbers of Asian or Pacific Islanders and American Indians who report being Hispanic, we include African Americans and whites who report being Hispanic in the Hispanic category only and include Hispanic Asian or Pacific Islanders in the Asian or Pacific Islander category and Hispanic American Indians in the American Indian category.

⁹ Data on the implementation date for TANF is available from the Assistant Secretary for Planning and Evaluation website: http://aspe.hhs.gov/hsp/Waiver-Policies99/policy_CEA.htm.

[Table 1 about here]

The mean teenage fertility rate from 1992 to 1999 was 54.4 births per 1,000 females aged 15 to 19. The fertility rate was higher among old teenagers compared to young teenagers: 85.9 versus 33.6. Table 1 also shows the differences in fertility rates by racial and ethnic groups. Teens who reported being Asian or Pacific Islander had the lowest fertility rate at 24.1 births per 1,000 teens, followed by white teens at 38 births per 1,000. American Indian teens had a fertility rate of 70.3, while African Americans and Hispanics had rates of approximately 96.5 births per 1,000 teenage females.

Empirical Strategy

To determine if teen fertility rates declined post-PRWORA, we estimate the following fixed effects weighted least squares (FE WLS) model:

$$(1) \quad \ln(\text{TFR})_{arst} = \alpha_0 + \alpha_1 \text{WR}_{st} + \alpha_2 \text{YT}_a + \alpha_3 \text{WR}_{st} * \text{YT}_a + \alpha_4 \mathbf{X}_{s(t-1)} + \alpha_5 \mathbf{A}_{st} + \alpha_r + \alpha_s + \epsilon_{arst},$$

where the subscript a represents age (15 through 19); the subscript r represents race/ethnicity; the subscript s represents the state; the subscript t represents the year; the variable TFR is the teen fertility rate; WR is an indicator for implementation of the minor parent rules (welfare reform);¹⁰ YT is an indicator equal to one for teenagers less than 18; \mathbf{X} is a vector that includes the natural logarithm of welfare benefits and the state unemployment rate; \mathbf{A} is a vector that includes an indicator if the state required parental consent for teenagers younger than 18 to receive an abortion and another equal to one if the state had a mandatory delay for abortion; α_r is a vector of indicators for African American, Asian or Pacific Islander, American Indian, and Hispanic, white is the omitted

¹⁰ We cannot separately identify the effects of the stay-in-school rule and the living arrangement rule.

category; α_s is a fixed state effect; and ϵ is a stochastic error term. We lagged the variables in \mathbf{X} one year to account for the gestation period for births. We weight all of the models by the age*race/ethnicity*state*year population of females. Further, we report Huber-White standard errors and correct for the intra-state dependence of observations.

From a high of 61.8 births per 1,000 teenagers in 1991, teenage fertility rates in the U.S. fell every year thereafter reaching a low of 48.8 births in 1999 (Hamilton, Sutton, and Ventura 2004). Given this downward trend in teenage fertility rates during the 1990s, we expect that the estimate of $\alpha_1 < 0$. We also know that young teens have lower teen birth rates compared to old teens (Hamilton, Sutton, and Ventura 2004); hence, we expect $\alpha_2 < 0$. If our hypothesis is correct, however, we anticipate a differential decline in teen birth rates for young women less than 18 after welfare reform, i.e., $\alpha_3 < 0$. Since our dependent variable is the natural logarithm of the teenage birth rate (births per 1,000 women), the coefficient, α_1 , indicates the percent difference in the birth rate for older teens before and after welfare reform while $(\alpha_1 + \alpha_3)$ indicates the percent difference in the birth rate for young teens. The coefficient, α_3 , indicates the percentage point difference between the percent change in the birth rate for young teens and the percent change in the birth rate for older teens.

Model 1 suffers from a variety of shortcomings. Obviously, any omitted variables that are correlated to teen fertility rates and age in the post-welfare reform period present the possibility of biasing our estimate of α_3 . The state fixed effects should capture any unobserved variables that are constant over time within a given state reducing the bias that omitted variables potentially produce. Unobserved factors that are correlated to age

and teenage fertility rates that vary over time remain a potential source of omitted variable bias.

In addition, it is highly unlikely that the average difference in teen fertility rates in the pre- compared to the post-welfare reform period was entirely due to PRWORA given that the teen birth rates were declining prior to the implementation of the program. It may be the case that the pattern in teen fertility rates across age groups was trending downward, but that this trend was more pronounced for young teens. Thus, when one observes average teen fertility rates before and after welfare reform and finds a larger gap between the rates for young and old teens in the post-reform period compared to the pre-reform period, one might be incorrect to conclude that the difference is due to welfare reform.

To allow for this alternative, we also estimate the following model:

$$(2) \quad \ln(\text{TFR})_{\text{arst}} = \beta_0 + \beta_1 \mathbf{T}_{\text{st}} + \beta_2 \text{YT}_a + \beta_3 \mathbf{T}_{\text{st}} * \text{YT}_a + \beta_4 \mathbf{X}_{\text{s}(t-1)} + \beta_5 \mathbf{A}_{\text{st}} + \beta_r + \beta_s + \eta_{\text{arst}}.$$

In Model 2, \mathbf{T} represents a vector of seven indicator variables representing the number of years that the minor parent provisions have been in place in that state (indicators from one to seven years). The coefficients for \mathbf{T} should be interpreted in contrast to the pre-reform period when the policies were not in place. Since we interact \mathbf{T} with an indicator for teens younger than 18, the coefficients in the vector, β_1 , represent the time trend in log teen fertility rates for females older than 17 following welfare reform. The coefficient vector β_3 , therefore, represents the percentage point difference in the time trend in teen fertility rates for young teens compared to old teens post-reform. If the coefficients in β_3 are statistically significant, then we interpret them to suggest that there are differences in fertility for young teens, relative to old teens, in the respective years. Further, one may

expect to find that the differences begin small as minors become aware of these provisions and increase over time eventually leveling out as all potential teen parents become aware of the policies.

Economic theory also suggests that since 15-year-old mothers in the welfare reform period lose benefits over a longer period of time than 17-year-olds, that these policies should have larger effects for teens the younger they are. To test this hypothesis we estimate the following model:

$$(3) \quad \ln(\text{TFR})_{\text{arst}} = \gamma_0 + \gamma_1 \mathbf{T}_{\text{st}} + \gamma_2 \mathbf{Age}_a + \gamma_3 \mathbf{T}_{\text{st}} * \mathbf{Age}_a + \gamma_4 \mathbf{X}_{\text{s}(t-1)} + \beta_5 \mathbf{A}_{\text{st}} + \beta_r + \beta_s + u_{\text{arst}},$$

where \mathbf{Age} is a vector of age indicator variables (19 is the omitted age) and the other variables are defined as before. We expect that the coefficient for $\gamma_3 < 0$ and will be larger in magnitude for the indicators for the younger ages.

Results

In the Model 1 of Table 2, estimate the importance of welfare reform on teenage fertility. Results from this model show a 5.2 percent reduction in teen fertility rates post-welfare reform.¹¹ It also suggests that young teenagers have fertility rates that are 67.4 percent lower than the rates for older teenagers. The coefficients for the racial and ethnic variables are as expected: African Americans, Hispanics, and American Indians have higher teenage fertility rates than whites, while Asian or Pacific Islanders have lower teenage fertility rates. State unemployment rates and welfare benefits are positively related to welfare benefits. Neither of the coefficients for the abortion variables is statistically significant.

[Insert Table 2 here]

¹¹ We calculate proportionate changes in teen fertility rates as $\exp[\alpha] - 1$.

In Model 2 of Table 2, we include the interaction between young teens and the post-welfare reform indicator variable (equation 1). This model suggests that fertility rates fell 9.2 percentage points more for young teens than for old teens and that this difference is statistically significant.¹²

Because the mean difference in log teen fertility rates pre- compared to post-reform are due to factors other than the changes in the welfare rules, we report results of equation 2 (which includes a series of indicator variables for the length of time since welfare was reformed in the state) in Table 3.¹³ We anticipate a nonlinear response to the time since welfare was reformed with the initial response being much smaller in magnitude, growing as teens become more familiar with the policy and then potentially leveling out. Of course, a priori, it is difficult to know if we have data of a sufficient duration to observe such a response.

In the first column we show that teen fertility rates appear to decline in percentage terms the longer the duration since welfare reform. Compared to the pre-reform period, the rates were approximately 4.4 percent lower after one year. By year four, the rates were 7.7 percent lower, and by the seventh year, the rates were approximately 11.3 percent lower. The decline in the rates over time conforms to our a priori expectations that the differences would grow over time.

[Insert Table 3 about here]

In Model 2, we interact the duration measures with the young teen indicator to determine if there is a different relationship for the young versus the old teenagers. This

¹² $-0.092 = [\exp(0.005 - 0.096) - 1] - [\exp(0.005) - 1]$

¹³ Since the earliest that teenagers could respond to welfare reforms is nine months after the implementation date, then we consider the year after implementation the first year post-reform. The indicators range from zero years (pre-welfare reform) to as high as seven years (Michigan was the only state to have seven post-reform years in our data).

model suggests that the declines post-reform are due almost entirely to the changes among the young teens relative to the old teens. The coefficients for the duration indicators (which represent the trend for old teens) are all small in size and statistically insignificant once we add the interactions. The coefficients for the interactions (which represent the difference in percentage changes in teen fertility rates for young teens) are all statistically significant. Further, the magnitude of the coefficients increases as the time since reform increases. For instance, for old teens the change in the fertility rate in the first year post-welfare reform was 0.5 percent lower. For young teens, however, the fertility rates in the first year post-reform were 6.7 percent lower, a difference of 6.2 percentage points. In the fourth post-reform year, the teen fertility rates among old teens was 0.2 percent higher, while the rates for young teens was 12.8 percent lower, a difference of 13 percentage points. Finally, in year 7, teen fertility rates for old teens were 0.9 percent higher and for young teens the difference was 18.8 percent lower, a difference of 19.7 percentage points.

To further test our hypothesis, in Table 4, we report coefficient estimates separating post-reform effects by the age of the teenagers. As explained earlier, relative to old teens, young teens are disproportionately harmed by the minor parent provisions. One might expect greater reductions in the log teen fertility rates for young teenagers relative to old teenagers. Model 1 shows, as expected, larger negative coefficients for younger ages (the omitted category is age 19). In Model 2, we interact the age indicators and a continuous measure of the number of years since welfare reform. The coefficients for all of the interactions are statistically significant, and results from this model suggest that the younger one is, the stronger the impact of welfare reforms. For instance, for 15-

year-olds, each additional post-welfare reform leads to a 5.4 percent reduction in teen fertility rates. In comparison, for 18-year-olds, the difference is a 0.8 percent *increase* in fertility rates.

[Table 4 about here]

If the teen mothers under the age of 18 are married, then they are not subject to the minor parent rules. Table 5 reports the percentage of births to unmarried teens by age, race, and ethnicity as reported by Ventura and Bachrach (2000) in 1992, 1996, and 1999. In the pre-reform era (represented by 1992), nearly all African American teenage births to young mothers were nonmarital. In comparison, between 68 and 82 percent of white births were nonmarital and between 65 and 76 percent of Hispanic births were nonmarital. Thus, a priori, we would expect welfare reform to have a greater impact on African Americans than non-Hispanic whites and Hispanics. Other differences across individuals, potentially correlated with race/ethnicity, also might yield differential effects of the minor parent rules on fertility. Thus, in Tables 6, 7, and 8, we repeat our analyses stratifying by race/ethnicity.

[Table 5 about here]

In Table 6, we report results from a FE WLS model of the difference in teen fertility rates for young teens compared to old teens by race and ethnicity (i.e., the Model 2 specification from Table 2). Table 6 shows a strong welfare reform effect for young African American teenagers. For old teens, the average teen fertility rate post-reform increased 3.4 percent, while the average teen fertility rate for young teens post-reform decreased 10.1 percent a difference of 13.5 percentage points. The difference was smaller for Hispanics, 2.7 percentage points, but statistically significant. For whites, the

difference was 8.8 percentage points and statistically significant. The percent difference is statistically insignificant for both Asian or Pacific Islanders and for American Indians. Interestingly, these results show a pattern that will persist throughout the remainder of this paper: the decline in fertility among young teens seems to owe primarily to changes among African Americans and whites.

[Table 6 about here]

In Table 7, we report results from Model 2 in Table 3 by racial and ethnic groups. A couple of findings are noteworthy. First, few of the post-reform year measures are statistically significant for older teens regardless of race. However, the pattern in the coefficients for the post-reform year indicators for both African Americans and whites are similar to that reported for the pooled sample. The coefficients for these two groups are statistically significant for every year post-reform and the magnitude of the coefficients are increasing during the post-reform period.

[Table 7 about here]

In Table 8, we report results from the Model 2 specification in Table 4 by race and ethnicity. Much like the results found in Tables 6 and 7, the results suggest that African Americans and whites are driving the relationship observed in the pooled data. All else equal, for 19-year-old African Americans, each year post-welfare reform is associated with a 2 percent increase in teen fertility rates. In comparison, among African American 15-year-olds, each year post-reform is associated with a 6.1 percent decline in teen fertility rates, a difference of 8.1 percentage points. Each age from 15 to 19 has a smaller negative coefficient. White teenagers follow exactly the same pattern, although the coefficient estimates are different. All else equal, for 19-year-old white teens, each

year post-reform is associated with a 0.1 percent increase in teen fertility rates. In contrast, for 15-year-olds each year is associated with a 5.1 percent decline, a 5.2 percentage point difference.

[Table 8 about here]

Conclusions

Much has been made of the major reforms of PRWORA: the introduction of work requirements, time limits, and the shift to a block grant with the creation of the TANF program. Another striking feature of the TANF program is that, unlike the AFDC program that preceded it, it made explicit the goal to reshape the family structure of low-income families. Theoretically, all of these reforms should have reduced the benefit of bearing a child out-of-wedlock for low-income women. Teenagers under the age of 18, however, were singled out in particular with PRWORA: Congress included a living arrangements rule and a stay-in-school rule requiring that teenagers under the age of 18 reside with a parent or other adult guardian and that they must be enrolled in high school (or have graduated) to receive welfare benefits. Thus, we hypothesized that in addition to the reduction in the probability of birth caused by the changes in TANF generally, young teens, in particular, should be less likely to bear children.

Using a fixed effects weighted least squares model our results support this hypothesis. We find that young teens had fertility rates that were 9.2 percentage points lower than older teens post-reform. Further, it appears that these changes can be attributed primarily to African American and white teens. This conclusion was supported in a subsequent analysis in which we found that the difference in log teen fertility rates

between young and old teens increases the longer the duration since welfare reform. Further, we find that these effects are strongest for the youngest teenagers which, again, comports with economic theory.

Of course, this evidence is not without limitations. Optimally, we would investigate the changes in fertility among 15- to 17-year-olds due to welfare reform in a high-risk or “treatment” group compared to a low-risk or “control” group. In this case, we would more accurately follow a quasi-experimental design. In the analyses we present, implicitly we assume that the trends for 15- to 17-year-olds would follow the same pattern as the trend for 18- and 19-year-olds, which may not be true.

We also can not be certain that the changes we identify are the results of the minor parent provisions. Any other changes that may have affected the fertility of minors that occurred simultaneously and were not controlled for will bias our estimates. For example, according to a survey by the American Public Human Services Association in 1999, 46 states used TANF money for teen pregnancy prevention programs (Levin-Epstein and Hutchins (2003); please see Cohen and Greenberg (2000) for a list of examples).¹⁴ Perhaps, then, it may be more accurate to consider our findings the result of welfare reform generally rather than due to the living arrangement rule and the stay-in-school rule exclusively. In fact, any change that may have impacted the fertility of teens whether associated with welfare reform or not but that occurred coincident with reform calls interpreting these results as a reform effect into question.

We defined welfare reform from the date of the implementation of the changes. There is some evidence that teens may have reacted to the public debate over reform or

¹⁴ In addition, Congress authorized \$50 million annually for up to five years for block grants to states to provide abstinence education (Mathematica Policy Research 2002). The funding was not available until 1998 so if these abstinence education programs had any impact it was likely after our observed data.

perhaps the authorization of these reforms rather than changing behaviors once the reforms were implemented (Horvath and Peters 1999). If true, then our results may be conservative estimates of the true impact of welfare reform.

In sum, our evidence suggests that changes in welfare can influence the fertility behavior of young teens, at least relative to that of older teens. While subject to a number of limitations, this study supports the idea that TANF can influence the family life of low-income families. Future research should consider the potential teenage fertility effects of the TANF program.

Data Appendix

Race/ethnicity-age-specific fertility rate: total number of births to women of a given age and race/ethnicity divided by the total population of females of the same age and race/ethnicity in the state multiplied by 1,000; Source: Birth numbers were extracted from the National Center for Health Statistics Natality Data Series CD-ROMs. We use the *Bridged-Race 1990-1999 Intercensal Population Estimates (Single-year of Age Detail) for Calculating Vital Rates* downloaded from the U.S. Census Bureau Web page: <http://www.cdc.gov/nchs/about/major/dvs/popbridge/popbridge.htm>. Accessed March 2, 2005.

Age: a set of indicator variables from age 16 to age 19 (15 is omitted) equal to one for the observations with the comparable age-specific fertility rate, e. g., Age 16 equals one for age 16 fertility rates, zero for fifteen-, seventeen-, eighteen-, and nineteen-year-old birth rates.

African American: indicator variable equal to one for non-Hispanic African American fertility rates; Source: same as fertility rates.

Hispanic: indicator equal to one for Hispanic white and Hispanic African American; Source: same as fertility rates.

Asian or Pacific Islander: indicator equal to one for women who reported being non-Hispanic Asian or Pacific Islander or Hispanic Asian or Pacific Islander; Source: same as fertility rates.

American Indian: indicator equal to one for women who reported being non-Hispanic American Indian or Hispanic American Indian. This indicator also includes Aleuts and Eskimos; Source: same as fertility rates.

State Unemployment Rate: Source: U.S. Bureau of the Census' *Statistical Abstract of the United States*, various years.

Welfare Benefits: Maximum AFDC/TANF amount per month for a family of four, inflated to 1999 dollars using the CPI-U-X1; Source: Robert Moffitt's publicly available data: <http://www.econ.jhu.edu/People/Moffitt/DataSets.html>. Accessed January 15, 2002.

Parental Notification/Consent Laws: an indicator variable equal to one in the years a woman, 18 or younger, was required either to notify or obtain parental consent before obtaining an abortion; Source: Phillip Levine (2004).

Mandatory Delay: an indicator variable equal to one in the years a state required a mandatory waiting period before allowing an abortion; Source: Levine (2004).

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Table 1: Weighted Descriptive Statistics

	<i>Mean</i> <i>(Standard Deviation)</i>
Fertility rate	54.43 (43.60)
Fertility rate for 15- to 17-year-olds	33.62 (29.31)
Fertility rate for 18- to 19-year-olds	85.93 (42.79)
Fertility rate for non-Hispanic African Americans	96.24 (45.48)
Fertility rate for non-Hispanic whites	38.03 (27.25)
Fertility rate for Hispanic African Americans and whites	96.96 (49.47)
Fertility rate for Asian or Pacific Islanders	24.09 (21.05)
Fertility rate for American Indians	70.29 (52.53)
Age	16.99 (1.42)
African American	0.151 (0.358)
Hispanic	0.132 (0.338)
White	0.665 (0.472)
Asian of Pacific Islander	0.040 (0.195)
American Indian	0.012 (0.110)
Welfare Benefits (in 1999 dollars)	513.56 (199.56)
State Unemployment Rate	5.88 (1.50)
Parental Consent Law for Abortion	0.456 (0.498)
Mandatory Delay for Abortion	0.158 (0.365)

Notes: all statistics weighted by the age*race/ethnicity*state*year population

Table 2: Fixed Effects Weighted Least Squares Models of Teen Fertility Rates between 1992 and 1999

	Model 1	Model 2
Post-Welfare reform	-0.053** (0.010)	0.005 (0.012)
Teen 15-17	-1.120** (0.016)	-1.162** (0.017)
Post-Welfare reform* Teen 15-17		-0.096** (0.007)
African American	1.107** (0.060)	1.107** (0.060)
Hispanic	1.126** (0.076)	1.126** (0.076)
Asian or Pacific Islander	-0.480** (0.132)	-0.480** (0.132)
American Indian	0.613** (0.091)	0.613** (0.091)
State Unemp. Rate	0.029** (0.004)	0.029** (0.004)
Welfare Benefits (ln)	0.757** (0.134)	0.758** (0.134)
Parental Consent	0.009 (0.012)	0.009 (0.012)
Mandatory Delay	0.006 (0.015)	0.006 (0.015)
N	9631	9631

Notes: * $p < 0.05$; ** $p < 0.01$; outcome is $\ln(\text{TFR})$; Huber-White standard errors reported; standard errors are corrected for intra-state dependence; models include state fixed effects; all models weighted by the number of teens in the age*race/ethnicity* state*year; 569 cases are missing due to zero births in that age*race/ethnicity*state*year cell.

Table 3: Fixed Effects Weighted Least Squares Models of Teen Fertility Rates between 1992 and 1999 Using a Vector of Indicator Variables for Years Since Welfare Reform

	Model 1	Model 2
Teen 15-17	-1.120** (0.016)	-1.162** (0.017)
Post-reform – Year 1	-0.044** (0.011)	-0.005 (0.013)
Post-reform – Year 2	-0.066** (0.011)	-0.003 (0.013)
Post-reform – Year 3	-0.072** (0.019)	-0.014 (0.021)
Post-reform – Year 4	-0.080** (0.028)	0.002 (0.030)
Post-reform – Year 5	-0.089** (0.035)	0.034 (0.039)
Post-reform – Year 6	-0.069** (0.037)	0.031 (0.040)
Post-reform – Year 7	-0.120** (0.034)	0.009 (0.036)
Post-reform – Year 1 * Teen 15-17		-0.064** (0.007)
Post-reform – Year 2 * Teen 15-17		-0.104** (0.009)
Post-reform – Year 3 * Teen 15-17		-0.098** (0.015)
Post-reform – Year 4 * Teen 15-17		-0.139** (0.035)
Post-reform – Year 5 * Teen 15-17		-0.205** (0.051)
Post-reform – Year 6 * Teen 15-17		-0.169** (0.020)
Post-reform – Year 7 * Teen 15-17		-0.217** (0.017)
N	9631	9631

Notes: * $p < 0.05$; ** $p < 0.01$; outcome is $\ln(\text{TFR})$; Huber-White standard errors reported; standard errors are corrected for intra-state dependence; models include state fixed effects; all models weighted by the number of teens in the age*race/ethnicity*state*year; 569 cases are missing due to zero births in that age*race/ethnicity*state*year cell; controls include indicators for African American, Hispanic, Asian or Pacific Islander, American Indian, state unemployment rate, welfare benefits, parental consent laws, and mandatory delay.

Table 4: Fixed Effects Weighted Least Squares Models of Teen Fertility Rates between 1992 and 1999 Using Age Indicators

	Model 1	Model 2
Age 15	-2.085** (0.028)	-2.050** (0.030)
Age 16	-1.209** (0.018)	-1.182** (0.020)
Age 17	-0.631** (0.014)	-0.616** (0.015)
Age18	-0.232** (0.007)	-0.225** (0.007)
No. of years since welfare reform	-0.019* (0.008)	0.022** (0.007)
Age 15 * No. of years since welfare reform		-0.077** (0.015)
Age 16 * No. of years since welfare reform		-0.057** (0.012)
Age 17 * No. of years since welfare reform		-0.032** (0.007)
Age18 * No. of years since welfare reform		-0.014** (0.004)
N	9631	9631

Notes: * $p < 0.05$; ** $p < 0.01$; outcome is $\ln(\text{TFR})$; Huber-White standard errors reported; standard errors are corrected for intra-state dependence; models include state fixed effects; all models weighted by the number of teens in the age*race/ethnicity*state*year; 569 cases are missing due to zero births in that age*race/ethnicity*state*year cell; controls include indicators for African American, Hispanic, Asian or Pacific Islander, American Indian, post-welfare reform, state unemployment rate, welfare benefits, parental consent laws, and mandatory delay.

Table 5: Percent of Births to Unmarried Teens by Age, Race, and Hispanic origin, 1992, 1996, and 1997

	15	16	17	18	19	Total
White, Non- Hispanic						
1992	81.8	74.7	67.7	59.7	49.9	59.7
1996	89.5	83.8	77.6	69.3	59.1	69.3
1999 ^a		84.5			67.5	72.6
African American						
1992	97	96.1	94.5	92.3	88.9	92.6
1996	98.7	98	97.4	95.5	92	95.4
1999 ^a		98.3			93.7	95.5
Hispanic						
1992	76.2	71.2	65.2	60.6	54.4	61.9
1996	80.8	76.5	72.1	66.2	59.9	67.7
1999 ^a		81.8			67.3	72.9

Source: Ventura and Bachrach 2000, Table 4.

Notes: Data in 1999 were not reported separately by age. Instead, the rates were grouped 15-17 and 18-19. Rates were not reported for non-Hispanic African Americans, Asian or Pacific Islanders, or American Indians.

Table 6: Fixed Effects Weighted Least Squares Models of Teen Fertility Rates between 1992 and 1999, by Race/Ethnicity

	African American	Hispanic	White	Asian or Pacific Islander	American Indian
Post-welfare reform	0.033* (0.013)	-0.001 (0.014)	0.001 (0.014)	-0.020 (0.042)	-0.059 (0.043)
Teen 15-17	-0.763** (0.026)	-0.886** (0.026)	-1.318** (0.015)	-1.156** (0.075)	-1.097** (0.015)
Post-welfare reform * Teen 15- 17	-0.140** (0.013)	-0.026* (0.010)	-0.093** (0.009)	-0.031 (0.048)	-0.034 (0.026)
N	1959	1976	2036	1808	1852

Notes: * $p < 0.05$; ** $p < 0.01$; outcome is $\ln(\text{TFR})$; Huber-White standard errors reported; standard errors are corrected for intra-state dependence; models include state fixed effects; all models weighted by the number of teens in the age*race/ethnicity* state*year; 569 cases are missing due to zero births in that age*race/ethnicity*state*year cell; controls include state unemployment rate, welfare benefits, and indicators for parental consent and mandatory delay.

Table 7: Fixed Effects Weighted Least Squares Models of Teen Fertility Rates between 1992 and 1999 Using a Vector of Indicator Variables for Years Since Welfare Reform, by Race/Ethnicity

	African American	Hispanic	White	Asian/Pacific Islander	American Indian
Teen 15-17	-0.763** (0.026)	-0.886** (0.026)	-1.318** (0.015)	-1.156** (0.075)	-1.097** (0.016)
Post-reform – Year 1	0.022 (0.013)	0.001 (0.014)	-0.005 (0.016)	-0.019 (0.033)	-0.047 (0.038)
Post-reform – Year 2	0.029 (0.013)	0.001 (0.017)	-0.003 (0.017)	0.023 (0.028)	-0.091 (0.049)
Post-reform – Year 3	0.012 (0.020)	0.033 (0.025)	-0.021 (0.022)	-0.057 (0.103)	-0.091 (0.056)
Post-reform – Year 4	0.046 (0.042)	0.005 (0.047)	-0.034 (0.031)	-0.036 (0.142)	-0.226 (0.156)
Post-reform – Year 5	-0.001 (0.027)	0.007 (0.050)	0.002 (0.044)	-0.103 (0.141)	-0.139 (0.107)
Post-reform – Year 6	0.008 (0.028)	0.173** (0.050)	-0.016 (0.043)	0.159 (0.174)	-0.035 (0.111)
Post-reform – Year 7	0.024 (0.045)	0.015 (0.065)	-0.035 (0.036)	0.331* (0.151)	-0.046 (0.035)
Post-reform – Year 1 * Teen 15- 17	-0.103** (0.012)	-0.020 (0.011)	-0.070** (0.010)	-0.006 (0.041)	0.001 (0.028)
Post-reform – Year 2 * Teen 15- 17	-0.153** (0.011)	-0.050** (0.011)	-0.109** (0.012)	-0.094* (0.040)	-0.056 (0.033)
Post-reform – Year 3 * Teen 15- 17	-0.154** (0.031)	0.004 (0.031)	-0.090** (0.018)	-0.010 (0.120)	-0.054 (0.044)
Post-reform – Year 4 * Teen 15- 17	-0.182** (0.057)	0.020 (0.037)	-0.098** (0.039)	0.166 (0.177)	0.033 (0.073)
Post-reform – Year 5 * Teen 15- 17	-0.162** (0.042)	0.022 (0.030)	-0.155** (0.064)	0.069 (0.110)	0.024 (0.119)
Post-reform – Year 6 * Teen 15- 17	-0.164** (0.026)	-0.108 (0.048)	-0.101** (0.016)	0.209 (0.133)	-0.131 (0.077)
Post-reform – Year 7 * Teen 15- 17	-0.245** (0.026)	0.021 (0.026)	-0.172** (0.015)	0.134 (0.075)	-0.409** (0.015)
N	1959	1976	2036	1808	1852

Notes: * $p < 0.05$; ** $p < 0.01$; outcome is $\ln(\text{TFR})$; Huber-White standard errors reported; standard errors are corrected for intra-state dependence; models include state fixed effects; all models weighted by the number of teens in the age*race/ethnicity*state*year; 569 cases are missing due to zero births in that age*race/ethnicity*state*year cell; controls include an indicator for post-welfare reform, state unemployment rate, welfare benefits, parental consent laws, and mandatory delay.

Table 8: Fixed Effects Weighted Least Squares Models of Teen Fertility Rates between 1992 and 1999 Using Age Indicators

	African American	Hispanic	White	Asian/Pacific Islander	American Indian
Age 15	-1.343** (0.041)	-1.590** (0.036)	-2.298** (0.024)	-1.983** (0.136)	-1.863** (0.033)
Age 16	-0.747** (0.029)	-0.868** (0.030)	-1.337** (0.019)	-1.204** (0.106)	-1.056** (0.024)
Age 17	-0.378** (0.020)	-0.445** (0.018)	-0.698** (0.016)	-0.686** (0.050)	-0.581** (0.013)
Age18	-0.111** (0.007)	-0.162** (0.008)	-0.259** (0.009)	-0.274** (0.030)	-0.190** (0.013)
Years since welfare reform	0.020* (0.008)	0.011 (0.007)	0.001 (0.007)	-0.007 (0.030)	-0.028 (0.020)
Age 15 * Years since welfare reform	-0.083** (0.017)	-0.028** (0.008)	-0.053** (0.011)	-0.040 (0.056)	-0.032 (0.021)
Age 16 * Years since welfare reform	-0.066** (0.016)	-0.014 (0.008)	-0.042** (0.009)	-0.001 (0.039)	-0.026 (0.012)
Age 17 * Years since welfare reform	-0.046** (0.011)	-0.010* (0.004)	-0.025** (0.006)	0.028 (0.023)	-0.015 (0.012)
Age18 * Years since welfare reform	-0.025* (0.007)	-0.013* (0.005)	-0.010** (0.003)	0.015 (0.018)	-0.001 (0.013)
N	1959	1976	2036	1808	1852

Notes: * $p < 0.05$; ** $p < 0.01$; outcome is $\ln(\text{TFR})$; Huber-White standard errors reported; standard errors are corrected for intra-state dependence; models include state fixed effects; all models weighted by the number of teens in the age*race/ethnicity*state*year; 569 cases are missing due to zero births in that age*race/ethnicity*state*year cell; controls include an indicator for post-welfare reform, state unemployment rate, welfare benefits, parental consent laws, and mandatory delay.