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Children as Income-Producing Assets: The Case of Teen Illegitimacy and Government Transfers

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This paper develops a classical model of the teen fertility decision in the presence of public income transfers. The theoretical model predicts that welfare payments will encourage fertility, holding constant other economic opportunities, and that better economic opportunities will discourage fertility. Considering the possible simultaneity of illegitimacy rates and benefit levels, due to the collective choice process, we confirm the theoretical model's predictions with state-level data from 1980 through 1990. We find that including fixed effects in the regression to control for unobserved differences between states does not sufficiently control for endogeneity. After controlling for endogeneity, real welfare benefits are strongly and robustly related to teen illegitimacy. The point estimates of the elasticity with respect to changes in the illegitimacy rate are around +1.3 for white teens and +2.1 for black teens. Real wages for women with a high school education or less are negatively related to teen illegitimacy for white teens, with an elasticity of around −0.4. Finally, male wages appear to have little effect on the illegitimacy rate for white teens but appear negatively correlated with the illegitimacy rate for black teens in some model specifications.

1. Introduction

The assertion that economic considerations play a significant role in family formation and fertility decisions is neither new nor controversial. The observation that there is a systematic interplay between economic considerations and fertility dates at least to Malthus (1798). In his *Essay on Population* (1798) and *Summary View on Population* (1830), he provided a series of conjectures and empirical evidence in support of the view that agricultural productivity provided an overall restraint on the positive and negative influences on birth rates. Subsequently refined

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and debated, the classical theory of population was summarized by Blaug (1978) as the proposition that

> . . . the production of children, [is] not as a means of spending income on "consumer goods" to acquire satisfaction, but as a method of investment in "capital goods" for the sake of a future return. (Blaug 1978, p. 78)

While this classical view has been adequate for loosely explaining population dynamics in agrarian societies, the modern economic theory of the family, due mainly to Becker (1991), views children as primarily a consumption, rather than an investment, good.[1] Undoubtedly, across most of the range of the income distribution in industrialized economies, the consumption view of children is the more suitable and powerful explanation. However, for individuals in poverty, various public cash and in-kind transfers create a series of economic incentives which, as we shall develop below, make the childbearing decision equivalent to the Malthusian analysis that children are income-producing assets as well as sources of utility. In the modern welfare state, it is the transfer system, rather than agricultural production, that creates income-producing opportunities.

The growth in public transfers has been accompanied by a sizable empirical literature in sociology and economics on the interaction between various family formation decisions and the welfare state. In general, the empirical evidence, from studies that have used differences in welfare benefits across time and across states to test whether the welfare system encourages illegitimacy, is inconclusive.[2] One possible reason for the mixed results is that illegitimacy might affect per recipient benefits either directly, due to voters' concerns about illegitimacy, or indirectly, because it affects the size of the state's welfare population. Increases in the size of the welfare population increase the cost to voters of providing a given level of benefits and, therefore, might cause voters to reduce per recipient benefits. In either case, benefit levels and illegitimacy are codetermined.

Once we control for this endogeneity, we find strong evidence that welfare has a large and statistically robust effect on illegitimacy. We find transfer elasticities with regard to teen illegitimacy rates on the order of $+1.3$ and $+2.1$ for white and black teens, respectively. In addition, we find own wage elasticities with regard to teen illegitimacy rates on the order of -0.4 for white teens and that own wage elasticities are not significantly different from zero for black teens.

We focus specifically on the effect that welfare has on out-of-wedlock teen fertility for several reasons. The first is that, although few Aid to Families with Dependent Children (AFDC) households are headed by teenagers (only 3–4%), a much larger proportion are headed by women who were teen mothers (around 40%) (General Accounting Office 1994). A second point is that teen mothers tend to be less educated and spend more time on AFDC than other participants.[3] Finally, about half of unwed teen mothers become welfare recipients within two years of the birth of their first child (General Accounting Office 1994). Overall, unwed mothers

[1] Another interesting example of the investment view of children is the English Poor Laws in the 18th and 19th centuries. Boyer (1990) finds that child allowances, a common form of poor relief for able-bodied laborers, had a positive effect on birthrates.

[2] See Moffitt (1992) or Murray (1993).

[3] Women who enter the welfare system under the age of 22 at the time of their first spell spend an average total duration of 8.23 years on AFDC; women who are between 22 and 30 spend only 7.08 years; and women aged 31 to 40 spend only 5.15 years (Committee on Ways and Means 1992).

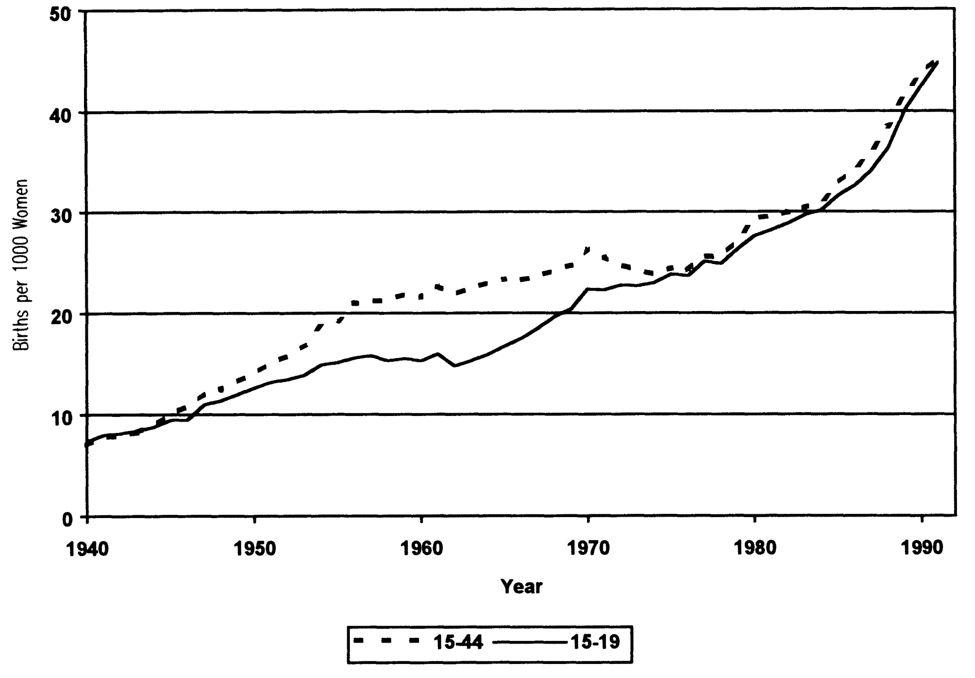

Figure 1. Illegitimacy Rates: 1940–1991

are likely to enter the AFDC program, and once this happens, they are in the program longer than older women.

The organization of the paper is as follows. Section 2 presents a brief review of empirical modeling considerations arising from the literature and some stylized facts. Section 3 develops and explores a formal economic model of the dependency decision. A young, fertile teenager is viewed as facing the choice to (i) complete her education and seek work or get married or (ii) have a child and thus gain access to AFDC, food stamps, Medicaid, and housing and energy assistance. Section 4 discusses the data collected to test the model and econometric modeling considerations. Section 5 presents and summarizes the empirical results, while section 6 concludes. Also, Appendices A and B contain, respectively, data definitions and sources and a discussion of sample size in other studies of teen illegitimacy that use individual-level data.

2. Some Stylized Facts on Illegitimacy and Welfare

Much of the recent concern about the effect of welfare benefits on illegitimacy has been stimulated by the large increase in births to unmarried women since the end of the Second World War. The number of illegitimate births per 1000 single women of childbearing age has nearly tripled over the past 40 years and has more than tripled among teens (see Figure 1).

It is commonly asserted that the increase in illegitimacy has been caused by welfare benefits encouraging, or at least allowing, single women to bear children out-of-wedlock. The objection to this conjecture is that, in real terms, the welfare benefits available to single mothers have not grown continuously over this period. One measure of the value of welfare benefits, the combined value of AFDC and food stamp payments to a family with no other income, grew slowly in the early 1960s and then faster through the mid-1970s.[4] However, since the mid-1970s, benefits have been flat or declining in real terms. Moffitt (1992) notes this stagnation makes it unlikely that changes in welfare benefits alone explain the rapid growth in illegitimacy.

This does not mean that benefits have played no role in the increase in illegitimacy. Even if the real value of welfare benefits has not been growing continuously, it might have been growing relative to other economic opportunities available to the young woman. In *A Treatise on the Family* (1991, p. 16), Becker observed that

> . . . my analysis of the marriage market indicates that the incentive to remain single depends upon income while single relative to income expected if married. The real wage rate of young male high school dropouts and the lowest quartile of graduates has dropped by more than 25% over the past 15 years and these young men may have become less attractive marriage partners for other reasons as well.

Welfare might have interacted with other variables to cause the rapid growth of illegitimacy, even if it is not the only, or even the main, contributor to the growth.

Looking at changes in illegitimacy and welfare benefits over time is one way of testing the relationship between the two. Another is to take advantage of the federal system where states set their own AFDC benefit levels.[5] As Murray (1993, p. 225) notes, this variation appears to ". . . provide a natural experiment for testing the proposition that welfare is linked to family breakup." If welfare were the primary cause of the increase in illegitimacy, then states with higher per-recipient benefits might have higher illegitimacy rates since women in those states might be more likely to give birth out-of-wedlock. Many studies have exploited differences in benefits across time or states, using discrete choice models to test whether welfare benefits affect the probability that an unmarried woman has an out-of-wedlock birth or to test the aggregate relation between benefit levels and the state's illegitimacy rate. However, empirical work exploiting these differences has not been conclusive. Some studies have found modest positive relations (e.g., Ozawa 1989; Caudill and Mixon 1993), many others have found mixed or statistically insignificant positive results (e.g., Duncan and Hoffman 1990; Lundberg and Plotnick 1990; Acs 1993), and others have even found negative correlations among their results (e.g., Ellwood and Bane 1985). In a recent paper, Rosenzweig (1995) finds that high AFDC benefit levels increase the probability that a woman will give birth out-of-wedlock before her 23rd birthday, especially for women who grew up in low-income households. Moffitt (1992) summarizes various studies written between 1982 and 1990 on the effects of welfare benefits and concludes that there is only "mixed evidence of an effect of the welfare system on illegitimacy." Murray (1993) and Acs (1993) examined other studies and reached the same conclusion.

However, differences in per-recipient benefits between states and across time are not the result of a natural experiment. Ellwood and Bane (1985) and others note that the state's benefit level is not set independently of the social and political structure of the state.[6] First, as noted by Ellwood and Bane (1985), omitted or perhaps unmeasurable state attributes might affect both the benefits the state offers single parents and encourage, or discourage, single motherhood. If these traits are omitted, the estimated coefficient on benefits will be inconsistent; this criticism

[4] Since food stamps were introduced in the late 1960s, this measure is usually extended prior to this by only counting AFDC benefits.

[5] This has led to benefits varying greatly between states—in 1991, the AFDC payment to a mother with one child and no other income varied between $120 per month in Mississippi and $694 per month in California. Including food stamps in the measure of total benefits reduces interstate differences, but differences still remain substantial.

[6] In the same way, variations in benefits across time may depend on changing social values or political structure.

applies equally to individual and aggregate studies. If these traits are (more or less) constant for a given state over the period studied or (more or less) constant across all states for a given time period, then state (or time) dummy variables in a regression analysis will control for them. However, illegitimacy rates might also affect benefit levels directly, either due to voters' concerns about illegitimacy or because they affect the size or composition of the state's welfare population. The public choice literature on how states set welfare benefits (e.g., Orr 1976) shows theoretically that the size of the welfare population (relative to the number of taxpayers) increases the price of per-recipient benefits for voters. If changes in the price of benefits to voters affect per-recipient benefits, then benefits and illegitimacy rates will be simultaneously determined.[7] Benefit levels might also affect the composition of the state's welfare population (e.g., changes in the relative number of divorced and never-married women), and this might also affect voters' preferred levels of benefits if voters are more sympathetic toward certain subgroups of recipients. If so, in either case, there is a simultaneity problem, where benefits and illegitimacy both affect the other. This is discussed further below.

3. A Model of Teen Fertility

Recently, the literature on fertility has concentrated on the interplay between child quantity and quality and on inter-generational transfers (e.g., Cigno 1986; Barro and Becker 1988; Becker 1991; Hanushek 1992). In these models, the number of children, child quality, and consumption goods enter the family's utility function, which is then maximized with respect to some budget constraints. Leisure is not usually included in the utility function.

In our simple model, a utility-maximizing woman faces a discrete choice between some combination of marriage and work on the one hand and welfare on the other. It is assumed that the woman has the number of children that she wants (i.e., there is no stochastic element to childbearing) and that her utility function is continuous and satisfies a nonsatiation condition. The price of consumption is normalized to one. Whichever choice she makes, she maximizes utility by choosing appropriate amounts of leisure, children, and market-consumption goods. For women who choose work and marriage, children are essentially a consumption good. However, for poor women on welfare, children also act as an income-producing asset. Table 1 contains variable definitions.

Children consume two different types of parental resources: (i) money—there is a financial cost, $p_b > 0$, associated with raising each child; and (ii) time—there is an additional time cost, $t_b > 0$, involved with raising each child. These costs are fixed and represent the minimum time and money investment the mother needs to bear and raise the child.[8] They are not substitutable—one cannot reduce the time commitment by increasing money expenditures and cannot reduce the financial commitment by increasing time expenditures.

[7] Empirical estimates that take into account the endogeneity of recipiency rates suggest that the elasticity of AFDC benefits with respect to the recipiency rate is probably not very large. For example, Ribar and Wilhelm (1996) and Shroder (1995) both conclude that their results indicate that the effects, if negative, are small and note that their results are sensitive to the estimation technique used.

[8] Results in this section are similar if the time cost of children is assumed to be an increasing, but possibly nonlinear, function of the number of children. To ensure in the welfare case that the woman's budget set is compact, it is necessary to make the additional assumption that either there is a physiological maximum on the number of children that the woman can have, that $g_2 < 0$, or that the time cost of an additional child is always greater than or equal to some δ that is strictly greater than zero.

Table 1. Variables in the Model

Variable	Interpretation
b	Children (babies)
c	Consumption
l	Leisure
I	Partner's income
t_b	Time cost of child
p_b	Money cost of child
w	Woman's wage
L	Woman's labor
m	Time endowment
$V(b, L)$	Cost of child care given b children and L hours of labor
g_1	Basic government welfare grant (guarantee)
g_2	Additional welfare grant per child

A woman who chooses marriage and work must divide her time between child rearing, work, and leisure and her income between child care if she works, child rearing, and consumption. Since her utility function has a nonsatiation property, the inequalities in the budget constraints are replaced with equalities. Since either partner's income I or hours spent working, L, could be zero, this case encompasses single mothers who work and married mothers who do not work, so that

$$\max U(b, c, l) \quad \text{such that } p_b^* b + c + V(b, L) = I + wL$$

$$l + t_b^* b + L = m.$$

The first constraint is the financial constraint. The woman spends her labor income wL and her partner's income I (assumed not to be a function of the number of children), on consumption, bearing and raising children, and child care. The second is her time constraint. She divides her time between leisure, child care, and labor. This constraint would hold in the same way for a single mother not receiving welfare, although her time cost for children may be different.

If the woman chooses welfare instead of work and marriage, it is assumed that she does not work and cannot marry, perhaps due to program requirements.[9] Since 1990, states operating AFDC programs have been required to operate an Aid to Families with Dependent Children for Unemployed Parents (AFDC-UP) program.[10] However, since the primary breadwinner must be unemployed to receive AFDC-UP payments, spousal income would still be zero.

Her maximization problem, if she chooses welfare, is therefore

$$\max U(b, c, l) \quad \text{such that } l + t_b b = m$$

$$p_b b + c = g(b),$$

where $g(b)$ is the welfare payment function, the money a woman receives from the state to support b children. The variables t_b and p_b are the time and money costs of having a child (not

[9] AFDC participants can work but, due to high child care costs or high marginal tax rates, over 90% do not. Committee on Ways and Means (1992) reports that, in 1990, only 8.2% of recipients earned any income.

[10] Committee on Ways and Means (1992). AFDC-UP provides aid to needy children in families where the primary breadwinner is unemployed.

necessarily the same as t_b^* and p_b^*, the time and money costs faced by a married woman). For simplicity, $g(b)$ is assumed to take the form

$$g(b) = g_1 + g_2 b \quad \text{for } b > 0$$
$$= 0 \qquad\qquad \text{for } b = 0.$$

Throughout, g_1 is referred to as the base welfare grant, the money all women on welfare receive, and g_2 is referred to as the additional grant for additional children. It is assumed that no consumption (of purchased goods) is extremely unattractive and that positive consumption is possible (i.e., for some b^*, $m - t_b b^* > 0$ and $g_1 + (g_2 - p_b)b^* > 0$). That is, if a woman chooses welfare, she will have at least one child and will, therefore, receive a strictly positive payment net of raising the child.[11] Note that g_2 may be either less than or greater than the minimum financial cost of having a child (p_b). Two results follow from nonsatiation.

(i) Increasing I, spousal income, or w, the woman's own wage, pushes the financial budget constraint for married/working women outward. Nonsatiation implies that the woman's utility must improve since the original bundle can still be obtained. Increasing w only weakly increases utility since, if the woman chooses not to work at either wage, her utility will remain unchanged.

(ii) Assuming no consumption is unacceptable, women on welfare will have at least one child ($g(0) = 0$). Therefore, increasing either g_1 or g_2 shifts her budget constraint outward and increases her utility.

Note that changes in benefits affect only welfare and changes in partner's income and own wages affect only work/marriage. Therefore, they unambiguously affect the relative attractiveness of the choices. Increases in welfare payments increase the attractiveness of welfare and hence should be associated with more women choosing welfare, while an increase in the income of potential partners or in the woman's own wage unambiguously makes marriage and work more attractive.

Even if a woman prefers more consumption to less, more children to fewer, and the additional welfare payment more than covers the minimum cost of an additional child, increases in the base welfare grant g_1 or the additional payment per child g_2 will not necessarily increase the number of children a mother on welfare will choose to have. The argument is similar to the argument that increasing wages does not immediately encourage the individual to work more. Having more children means less time is available for leisure. Since she cannot have more of both, which one she chooses depends on her preferences. In the same way, increasing I (partner's income) or w (the woman's own wage rate) will not necessarily cause a woman who chooses work/marriage to have more children either. This result does not rely on obscure utility functions where children are not a normal good. If $g_2 > 0$, the number of children is a decreasing function of g_1 with a Cobb–Douglas utility function.[12]

In summary, the theoretical model makes the following testable predictions.

(i) Increasing either the base welfare grant or the additional welfare grant per child increases utility from welfare and does not affect the utility from marriage and work. Hence, we would expect more women to choose welfare over work and marriage when welfare payments increase.

(ii) Increasing spousal income or the woman's own wage will increase (weakly in the case

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[11] For example, assuming that $u(b, 0, l) < u(b', c', l')$, where $b, l, b', l' \geq 0$ and $c' \geq 0$, will ensure this.

[12] Proofs of these assertions are available from the authors on request.

of own wage) utility from work and marriage and does not affect utility from welfare. Hence we would expect fewer women to choose welfare as spousal income or women's wages increase.

(iii) Both the base grant g_1 and the additional grant for extra children g_2 have ambiguous effects on the number of children that a woman on welfare chooses and it is possible that changes in the base and additional grants could have opposite effects. An increase in the additional grant g_2 is more likely to have a positive effect since it is more likely to encourage the woman to have additional children.[13] These points are important for two reasons. First, the illegitimacy rate is affected by both the number of women choosing welfare and the number of children women on welfare have. Therefore, benefit changes have an ambiguous theoretical effect on illegitimacy. Focusing on teens reduces this concern because most births to teens are first births. Because women must have at least one birth to qualify for AFDC, the effect on first births is unambiguous. Second, the empirical literature has often not distinguished between base and additional grants in regressions relating out-of-wedlock fertility to welfare payments. Going beyond the simple theoretical model presented in this section, g_1 may be more important when initially choosing welfare, especially if teens have high discount rates.

4. Data and Econometric Issues

Issues in Measuring Illegitimacy

The empirical estimation focuses on out-of-wedlock births to teenagers rather than out-of-wedlock births to women of all ages for three reasons. First, as noted earlier, women on welfare who give birth as teenagers tend to be on welfare for longer and tend to become more dependent on welfare than women who give birth when older. Second, we wish to focus on the choice between welfare and work or marriage rather than on the number of children that unmarried women have. The illegitimacy rate—defined as the number of out-of-wedlock births per 1000 women of childbearing age—could be affected by welfare in two ways. The base and additional welfare grants might affect both the number of women that give birth out-of-wedlock and the number of children the women have. This is a concern because, as noted in the theoretical section, welfare might affect these decisions differently (in both direction and magnitude). Since most births to teens are first births, this distinction is less important for teens than for older women.[14] Finally, we focus on teens because women of different ages might be affected by welfare differently. It is easier to interpret the coefficients when the population is homogeneous.

Specification Issues

The theoretical model suggests that the left-hand-side variable should be either the illegitimacy rate or the observation of an out-of-wedlock birth.[15] It also suggests several right-hand-

[13] Proof available from the authors on request.

[14] In 1989, 77% of out-of-wedlock births to teens were first births (83% for white teens and 69% for black teens), whereas for women over 20, only 36% were first births. The percentage for teens varied between 77 and 79% over the decade.

[15] The illegitimacy rate, defined as the number of out-of-wedlock births per 1000 women, is only affected by the women choosing to have out-of-wedlock births. The illegitimacy ratio, defined as the ratio of out-of-wedlock to in-wedlock births is also affected by the number of women who have in-wedlock births. Since changes in male and female wages might affect the childbearing decisions of married women, it would not be possible to disentangle these effects using illegitimacy ratios.

side variables that should be included in a reduced form model of the woman's decision. These include the value of the base and additional AFDC grants, the wage available to the woman if she does not have a child out-of-wedlock as a teen, the wage or income of the woman's prospective spouse if she does not have a child out-of-wedlock, and perhaps the availability of potential spouses. Beyond the issue of the appropriate dependent and independent variables, there are several conceptual and econometric issues.

(i) *Unobserved Differences Between States and Across Time.* One issue frequently discussed in the literature on the effects of welfare on family formation decisions is the inclusion of fixed effects.[16] Ellwood and Bane (1985) note that omitted variables, such as unobserved state-level social and political characteristics, might affect both the probability of out-of-wedlock births and the AFDC benefit the state offers. For example, Ellwood and Bane (1985) suggest Minnesota's Scandinavian tradition might encourage both strong family ties and generous welfare benefits. In addition to inherently unmeasurable differences in attitude, one can also think of other omitted variables (potentially measurable and potentially unmeasurable) that might be correlated with both benefit levels and the prevalence of out-of-wedlock births. If these variables are (approximately) constant over time (for each individual state) or (approximately) constant across states (for each time period), then time and state dummies will effectively remove them from the regression, allowing unbiased estimation of other parameters. If the variables are not constant, then, in general, fixed effects estimation will not be consistent. This problem becomes more likely as the period studied becomes longer. Omitted variables that are similar when comparing 1983 to 1984 may be far less similar when comparing 1960 to 1995.

(ii) *Endogeneity of AFDC Benefits.* AFDC benefit levels are the result of decisions made directly by state-level politicians, and therefore indirectly by voters, in each state. Bearing this in mind, there are at least two reasons that AFDC benefits might be endogenous. First, as noted above, there may be omitted societal variables that affect both the collective decision regarding benefit levels and the individual decision of whether to give birth out-of-wedlock. Second, teen pregnancy rates might affect the benefit level directly rather than through other omitted variables, resulting in a classic simultaneity problem. This might occur if some voters, concerned about teen pregnancy, believe teenagers have children out-of-wedlock to receive welfare payments. These voters, then, might try to cut benefits to discourage out-of-wedlock childbearing. In addition, out-of-wedlock births to teens might (at least in the long run) affect either the size or the composition of the welfare population. This might, in turn, affect voters' perceptions of welfare and the generosity of benefits. The public choice literature on state welfare policies shows, theoretically, that the size of the welfare population (relative to the number of taxpayers) is the fiscal price of per-recipient benefits to voters. The intuition is that, as the number of recipients increases, it becomes more costly to pay a given per-recipient benefit (Orr 1976). Hence, if the number of teens giving birth out-of-wedlock affects the size of the welfare population, there is a simultaneity problem. The composition of the welfare population might also be important. For example, voters may be more or less sympathetic toward widows or divorced recipients (when compared to never-married recipients). If changes in benefit levels change the relative numbers of women in each group (for example, if one group is more responsive than the other to benefits), endogeneity will be a concern.

These problems occur in studies using both individual and aggregate data, although they can be handled in different ways. It is clear that omitted variables, correlated with both benefit

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[16] See, for example, Jackson and Klerman (1994), Moffitt (1994), and Hoynes (1997).

levels and illegitimacy, are a concern in both types of studies. Therefore, including state and time dummies is likely to be important when using either aggregate or individual data. Endogeneity is an obvious concern for aggregate studies, such as this one, but could also be relevant for individual studies. If the error term (which may include omitted variables) is independently distributed across teens within the state, then endogeneity caused by the aggregate illegitimacy rate affecting benefit levels would not be a large concern. Any individual teen's decision would have only a small net effect on aggregate birth rates, and so the benefit level would be trivially correlated with the individual's error term. However, if there are omitted variables that are correlated across teens in the state (violating the assumption of independently distributed errors), the effect on aggregate illegitimacy rates might not be trivial. This makes endogeneity a concern whether the omitted variables affect benefit levels directly or not. If these omitted variables are constant for each state over the entire period (constant across states for each period), then including state (time) dummies will be sufficient. If not, then dummy variable estimation will not give consistent parameter estimates when using individual data.

Individual Versus Aggregate Data

In this study, we use aggregate state-level data rather than individual data. Although individual data have several advantages over aggregate data, there are some reasons to favor the latter.[17] The main problem with individual data is that relatively small samples, and the relative rarity of out-of-wedlock births to teens (especially white teens), make it difficult to estimate empirical models when state dummies are included in the regression.[18] For white teens, the main problem is that out-of-wedlock births are relatively rare. In a sample of 600 white, female teens from the National Longitudinal Survey of Youth (NLSY), only 44 births are observed in only 22 states (see Table B1 in Appendix B). This will obviously make it difficult to include 50 state dummies in the regression. If there are no births in the state, then including a state dummy removes all observations for that state from the regression, reducing variation in the AFDC variable and decreasing true sample size. For black teens, although birth events are more common, samples tend to be smaller and more geographically concentrated. This issue is discussed in greater detail in Appendix B.

An additional point when using individual data is that it is important to be careful when trying to control for the other economic opportunities available to the teen. In particular, including individual characteristics in the regression might not fully control for wage differences across states and for changes in wage structure across time. For example, if wages for high school graduates are lower in Mississippi than in California, then even controlling for education

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[17] For example, panel data that follow individuals across time, such as the NLSY (Plotnick 1990; Acs 1993; Lundberg and Plotnick 1994) and the Panel Study of Income Dynamics (PSID) (Duncan and Hoffman 1990), allow the researcher to control for a wide range of individual-specific, family background, and neighborhood effect variables (i.e., parental education levels, number of siblings, and birth order) in the estimation. In addition, they allow the researcher to address questions that cannot be easily addressed with aggregate data, such as the effect of childhood events (An, Haveman, and Wolfe 1993) or church attendance (Plotnick 1990).

[18] This problem is more severe when examining the effect of welfare benefits on out-of-wedlock births to teens than when studying other questions, such as whether or not welfare benefits encourage female headship in the population (see Moffit 1994; Hoynes 1997). For these more general questions, sample sizes tend to be larger and the events more common.

level (which is not observed for teens) would not fully capture wage differences.[19] The data used in this study indicate that wages at the bottom of the income distribution are highly correlated with benefit levels. In 1990, women's and men's wages (for persons with a high school education or less, aged between 21 and 35) have respective correlations of 0.77 and 0.75 with combined AFDC, food stamp, and Medicaid benefits for a family of two. Therefore, it seems important to at least include aggregate measures of wages in individual regressions to avoid potential bias.

Data

The data used are aggregate state-by-state data for 1980 through 1990 and are documented in Appendix A. The dependent variable is the illegitimacy rate. An important question is whether births to all teens or only births to unmarried teens is a better numerator. As Acs (1993) points out, the fertility of married teens may also be affected by changes in welfare benefit levels if married teens see welfare as insurance against divorce. However, it seems unlikely that changes in welfare benefits would affect married teens to the same extent as unmarried teens. Furthermore, the marriage decision itself may be influenced by the welfare benefit level. For this reason one should interpret the results in this paper carefully; the effect of government benefits and wages on illegitimacy may be primarily due to effects on marriage rather than on fertility. A positive coefficient on government benefits might not imply that welfare encourages more teens to get pregnant, just that fewer teens get married when they do become pregnant. Finally, we note that out-of-wedlock births among teenagers might be thought to be a greater social problem than in-wedlock births in terms of welfare dependency and outcomes for children.

The three main explanatory variables of interest are (i) the value of AFDC, food stamps, and Medicaid benefits for a family of two; (ii) the median weekly wage of women between 21 and 35, working full-time, with a high school diploma or less; and (iii) the median weekly wage of men between 21 and 35, working full-time, with a high school diploma or less.[20]

The effect of the additional grant (the difference between the value of the benefit package for a family of two and a family of four) is also tested. The wage variables represent the value of work and marriage to women at the lower end of the income distribution. A potential problem with these regressors is that the three variables are highly correlated across states. This might make it difficult to interpret the effects of each variable separately with a high degree of confidence. An additional included variable is the incarceration rate. This is intended to control for

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[19] Another point is that controlling for the teen's (and her potential partner's) other economic opportunities is harder for teens than for older women because available proxies are weaker. The most reasonable proxy for individual wages, completed education level, is not observed for teens. At age 15, most teenagers, whether they eventually complete only high school or if they go on to get a doctorate, will have similar levels of education. Duncan and Hoffman (1990) estimate predicted income at age 26 for a sample of black teenagers from the PSID. The women were teens between 1968 and 1985. They find that income at age 26 has a strong and statistically significant effect on childbearing and that the effect of AFDC is weak and statistically insignificant. However, as they note, most of the variables they use to predict age 26 income if the woman did not give birth as a teen could be included in the birth regression, resulting in linear dependence among regressors. Additionally, they do not include state or time dummies in either regression. If, as noted below, AFDC benefits are correlated with wages across states or time, this could be problematic.

[20] Results using just the combined food stamp and AFDC guarantee are broadly similar to the results presented here. See footnote 51 for a full description of the differences between the two sets of results. Analogous tables with the AFDC and food stamp guarantee included in place of the AFDC, food stamp, and Medicaid guarantee are available from the authors on request.

the size of the pool of marriageable men (Garfinkel and McLanahan 1986; Wilson 1987).[21] Recent increases in illegitimacy may be partially due to a decline in the number of men available as potential marriage partners. The incarceration rate is likely to be correlated with other factors, such as high drug use and high mortality rates, also related to this concern. However, it is also possible that this variable may simply pick up increased juvenile delinquency. Because of this, a precise interpretation of the coefficient is difficult. In conclusion, the predictions from the theoretical model are that wages will be negatively correlated with the illegitimacy rate among teens, welfare benefits will be positively correlated with the illegitimacy rate among teens, and the incarceration rate will be positively correlated with the illegitimacy rate among teens.

Additional variables are included as controls. The availability of abortions is proxied by the percentage of counties in the state with an abortion provider, an admittedly imprecise proxy. It is plausible that easy access to abortion may reduce the number of births to unmarried teens. Although easier access to abortion may encourage sexual activity among teens, it seems reasonable to suppose that only teens who would choose to have an abortion if pregnant would be encouraged to become sexually active. However, Akerlof, Yellen, and Katz (1996) note, in the context of a theoretical model, that it is possible that increased access to abortion might also increase sexual activity among individuals who would not obtain abortions if pregnant.[22] They write

> Before the technology shock (the introduction of abortion or contraception) abstinence would be the norm for all women. After the technology shock those women who would use contraception or would be willing to obtain an abortion in the event of pregnancy or both engage in premarital sexual activity. However, those women who are not willing to use contraception or obtain an abortion will also engage in sexual activity, since they fear that if they abstain their partners would seek satisfaction elsewhere. The advent of contraception and abortion used by others may result in an unwanted increase in sexual participation for those who reject the new technology. (p. 296)

The unemployment rate and female unemployment rate are included as measures of the state of the labor market. Finally, the percentage of the population living in metropolitan areas and the infant mortality rate, a common variable in studies of fertility (see Shields and Tracy 1986), are also included.[23]

Instruments

An immediate problem is finding variables that can serve as instruments for the AFDC benefit level. This is difficult because many variables that would seem likely to affect the AFDC benefit level might also affect the illegitimacy rate. For example, demographic variables, such

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[21] We also test an additional variable, the ratio of male to female teens aged 15–19, to control for the pool of marriageable men. Results for the benefit and wage variables are similar in terms of both size and statistical significance when this variable is included in the regression. For black teens, the coefficient on this variable is consistently statistically insignificant, while for white teens it has a counter intuitive positive sign in most regressions (i.e., the number of male teens per female teen is positively correlated with the number of births).

[22] Improved access to birth control might have an ambiguous effect on out-of-wedlock childbearing even if only teens who would use birth control are encouraged to become sexually active. Access to birth control might encourage teens who would give birth if they became pregnant to become sexually active, as well as teens who would have an abortion. As a result, the aggregate effect depends on the number of teens (who would not have an abortion if pregnant) who switch to more effective birth control methods and the number who switch from abstinence to less effective forms of birth control.

[23] It has been suggested that infant mortality might be a consequence, rather than a cause, of teen fertility (see, e.g., Pampel and Pillai 1986; Cramer 1987; Bennet 1992). However, results in this study are similar in terms of size and statistical significance whether this variable is included or not.

as the share of the population living in urban areas, might affect the AFDC benefit level but also might affect the illegitimacy rate. Because of this, it seems important to have more instruments than endogenous variables so that the overidentifying restrictions can be tested. Another reason that we want more than one instrument is that we would like to be able to separately test the effects of the base and additional grants.

The public choice literature on welfare benefits suggests several possible instruments for AFDC benefits.[24] The main variables used in the public choice literature are income of the median voter and the price of benefits. The price of benefits is the recipiency rate multiplied by the state's share of costs (1 − federal matching rate). Even if benefit levels have no incentive effects (i.e., benefit levels do not directly affect the behavior of potential recipients), more people will qualify for AFDC as benefit levels increase, making the recipiency rate endogenous. Therefore, we use the federal matching rate rather than the recipiency rate multiplied by the federal matching rate as an instrument. Federal matching varies between 50% and 83% of the total benefit and depends on the state's per capita income.[25] Federal matching affects the cost of benefits to the state and therefore is likely to affect the state's benefit level but would not directly affect the choice of the teen. Median voter income (proxied by per capita income) might be an acceptable instrument if teens are mainly affected by movements of income at the bottom of the income distribution. Another plausible instrument is the percent of the population that is over 65. This might affect the distribution of public funds but should have little effect on the illegitimacy rate. Using all three instruments, we can test the overidentifying assumptions.[26]

A potential problem is that per capita income, which is correlated with the wage measures, might be correlated with the benefit level even after controlling for wages at the bottom of the income distribution (e.g., perhaps due to nonlinear relations between wages and illegitimacy rates). Further, since the matching rate is a function of state per capita income (relative to national per capita income), this would make this instrument endogenous also. Although tests of overidentifying assumptions are included to ease concerns, we also present results using the percent of the population that is over 65 as the sole instrument and including per capita income as an independent variable.

5. Empirical Results

Econometric Modeling

The basic model is

$$\text{illegitimacy rate}_{it} = \alpha_i + \gamma_t + \beta' x_{it} + \epsilon_{it},$$

where i indexes state and t indexes time. There is an observation for each state for each year

[24] See, for example, Orr (1976) or Moffitt (1990b).

[25] In practice, the highest federal matching rate was 78.85%.

[26] Besley and Case (1994) suggest using political variables, such as the composition of the state legislature, the governor's party, and whether the governor can run for re-election, as instruments for policy variables. They note that these variables make it quite clear where the variation in policy is coming from. However, it is not obvious that these variables are appropriate instruments for welfare benefits. If voters believe that welfare is an important issue, welfare and illegitimacy might directly affect the outcome of elections. However, as discussed in Besley and Case (1994), in this case also, the abundance of political variables means overidentifying assumptions can be tested. In practice, we found that state-level political variables performed very poorly as instruments. In particular, tests of overidentifying restrictions rejected the null hypothesis that the instruments were valid, and the instruments were only weakly correlated with benefit levels.

Table 2. Results from Empirical Model by Regression Type

	(1) OLS	(2) OLS	(3) OLS	(4) OLS
State Dummies	No	No	Yes	Yes
Time Dummies	No	No	Yes	Yes
Dependent Variable	Illegitimacy rate for white teens	Illegitimacy rate for black teens	Illegitimacy rate for white teens	Illegitimacy rate for black teens
Instruments	None	None	None	None
AFDC, Food Stamps, and Medicaid	−0.0001	−0.0367[d]	0.0055[c]	−0.0192
	(−0.03)	(−2.66)	(1.80)	(−1.00)
Female Wage	0.0036	0.0818	−0.0113	0.1107[d]
	(0.22)	(1.23)	(−1.41)	(2.21)
Male Wage	−0.0188[d]	−0.0039	0.0027	0.0018
	(−2.45)	(−0.12)	(0.68)	(0.08)
Incarceration Rate	0.0230[d]	0.0401[d]	0.0113[d]	0.0477[c]
	(7.26)	(3.09)	(2.86)	(1.93)
Abortion	−1.1260	−51.831[d]	−4.1016[c]	0.1702
	(−1.00)	(−11.18)	(−1.87)	(0.01)
Unemployment Rate	1.3948[d]	1.3193	0.1624	−2.4723[c]
	(3.65)	(0.84)	(0.80)	(−1.95)
Unemployment Rate for Women	−1.3192[d]	−3.4595[d]	−0.2572	1.6915
	(−3.09)	(−1.98)	(−1.24)	(1.31)
Infant Mortality Rate	−2.1827[d]	0.7063	0.0338	1.2703
	(−11.31)	(0.89)	(0.24)	(1.47)
Metropolitan Population	0.0088	0.6096[d]	−0.0543	−0.1453
	(0.64)	(10.84)	(−0.89)	(−0.38)
Per Capita Income				
Overidentifying Restrictions χ^2 (2)	—	—	—	—
R^2	0.3599	0.3690	0.9308	0.8408
Number of Observations	535	535	535	535

t-statistic in parentheses.
[a] Public choice instruments are per capita income, percentage of the population over 65, and the Medicaid matching rate.
[b] AFDC program instruments are the difference between the percent of the state population and state AFDC recipient population who are black and administrative cost per AFDC family.
[c] Significant at the 10% level.
[d] Significant at the 5% level.

from 1980 through 1990.[27] The error term consists of (i) α, a state effect; (ii) γ, a time effect; and (iii) ϵ, the remaining individual error for that observation. The model is estimated with standard panel data techniques. Previous research has noted that fertility outcomes for black and white teenagers are different, and so this study estimates separate equations for black and white teenagers.[28]

[27] Arizona is omitted because it did not have a comparable Medicaid program over this period. Alaska and Hawaii, which have separate food stamp guarantees and food stamp income disregards, are missing data for 1980 and 1981.

[28] See Plotnick (1990). In addition, since some states have few black residents, results excluding states with few black residents are available from the authors on request. Fourteen states with less than 2000 black teens aged between 15

Table 2. Extended

(5) 2SLS	(6) 2SLS	(7) 2SLS	(8) 2SLS	(9) 2SLS	(10) 2SLS
Yes	Yes	Yes	Yes	Yes	Yes
Yes	Yes	Yes	Yes	Yes	Yes
Illegitimacy rate for white teens	Illegitimacy rate for black teens	Illegitimacy rate for white teens	Illegitimacy rate for black teens	Illegitimacy rate for white teens	Illegitimacy rate for black teens
Public choice instruments[a]	Public choice instruments[a]	Percent of population over 65	Percent of population over 65	AFDC program instruments[b]	AFDC program instruments[b]
0.0538[d]	0.3432[d]	0.0615[d]	0.3484[d]	0.1101[c]	0.0766
(2.95)	(2.79)	(2.51)	(2.23)	(1.95)	(0.41)
−0.0294[d]	−0.0251	−0.0321[d]	−0.0447	−0.0594[d]	0.0465
(−2.45)	(−0.31)	(−2.09)	(−0.46)	(−2.10)	(0.49)
−0.0054	−0.0591	−0.0066	−0.0768	−0.0125	0.0072
(−0.95)	(−1.54)	(−0.80)	(−1.46)	(−1.03)	(0.18)
0.0071	0.0159	0.0064	0.0164	0.0002	0.0402
(1.38)	(0.46)	(1.15)	(0.46)	(0.02)	(1.27)
−3.4365	5.1605	−3.3380	6.3913	−5.9441	−2.1645
(−1.26)	(0.28)	(−1.14)	(0.34)	(−1.26)	(−0.14)
−0.1950	−5.1537[d]	−0.2535	−5.0054[d]	−0.7895	−3.1909[c]
(−0.69)	(−2.71)	(−0.81)	(−2.51)	(−1.35)	(−1.63)
−0.0708	3.0903[c]	−0.0416	3.2161[c]	0.4056	1.9420
(−0.27)	(1.73)	(−0.14)	(1.73)	(0.82)	(1.18)
0.1761	2.3381[c]	0.1987	2.3815[c]	0.2262	0.7952
(0.98)	(1.94)	(1.01)	(1.90)	(0.77)	(0.81)
0.0672	0.7661	0.0865	0.7935	0.2786	−1.4129
(0.77)	(1.30)	(0.86)	(1.23)	(1.01)	(−1.54)
		−0.0001	0.0147		
		(−0.03)	(0.74)		
0.62	0.94	—	—	0.51	0.09
—	—	—	—	—	—
535	535	535	535	486	486

As a first exercise, mainly for comparison with later results, columns 1 and 2 in Table 2 show results from a simple OLS regression omitting state and time dummies and treating benefits as exogenous. These preliminary results do not support the theoretical model. For white teens, the coefficient on AFDC benefits has a theoretically incorrect, but statistically insignificant, negative sign (indicating high benefits are correlated with low rates of illegitimacy). The median weekly wage for women aged between 21 and 35, working full-time, with a high school education or less (referred to as the female wage) is statistically insignificant. The median weekly wage for men between 21 and 35, working full-time, with a high school diploma or less (referred to in the tables as the male wage) has the expected sign and is significant at a

and 19 are excluded from the regression for black teens in the reduced-sample estimates. However, we are unable to reject the null hypothesis that the estimated coefficients for the subsample with few black teens are different from the estimated coefficients from the estimation for the remaining states. Therefore, the results from the larger sample would appear preferable. We note in the text where results for the smaller sample differ significantly from results for the larger sample.

5% level. The results for black teenagers (column 2) are even less encouraging—only the male wage and the incarceration rate have the theoretically expected signs, and the male wage variable is statistically insignificant. Further, the benefit variable has a theoretically incorrect, and statistically significant, negative sign.

Columns 3 and 4 show results from basic two-way fixed effects or least squares dummy variable regressions, treating combined AFDC, food stamp, and Medicaid benefits as exogenous. For white teens, the coefficients on female wages and government benefits have the anticipated signs but are insignificant at a 5% level (although the coefficient on government benefits is significant at a 10% level). The coefficients are also quite modest—the implied elasticity of illegitimacy with respect to government benefits is 0.18 and the implied elasticities of female and male wages are −0.16 and 0.05, respectively.[29] For black teens, the coefficients on all the main variables have theoretically incorrect signs and the coefficient on female wages is statistically significant.

Columns 5 and 6 of Table 2 show results when state and time effects are included and the benefits variable is considered endogenous (using the public choice instruments—the AFDC matching rate, per capita income, and percentage of the population over 65).[30] For white teens, the coefficient on the benefit variable becomes large in absolute value, significant at a 5% level, and has the expected sign. The implied elasticity with respect to benefits (estimated at the means of all variables) is 1.81. The coefficient on female wages is also larger in absolute value and is significant at a 5% level. The implied elasticity is 0.41. The coefficient on male wages is insignificant at conventional levels, small in absolute value, but has the expected sign. The results for black teenagers are also closer to those predicted by theory. The coefficient on the benefits variable is positive and is significant at a 5% level, with an implied elasticity of 2.66. The coefficient on male wages is negative but is only significant at a 13% level. The coefficient on female wages is statistically insignificant, although it does have the expected sign.

Because these results vary greatly, a first question is which estimation technique is appropriate. For both OLS and 2SLS and for both black and white teenagers, the null hypothesis that the time and state dummy variables are jointly insignificant is rejected at a 1% level.[31] For white teens, the null hypothesis that the time dummies follow a simple trend is rejected in favor of the alternative hypothesis of time dummies. For black teenagers, the null hypothesis of a time trend can generally not be rejected in the 2SLS and generalized method of moments (GMM) regressions. However, the results are similar when a time trend, rather than time dummies, is included in terms of both size and statistical significance.[32] Overall, these results favor models with state and time dummies.

Broadly speaking, there are at least three criteria to consider when selecting instruments:

(i) Whether the moment restrictions imposed are valid. Tests of overidentifying assumptions might ease concerns regarding whether the instruments are exogenous.

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[29] Elasticities are evaluated at the means of the respective variables.

[30] Generalized method of moments results are similar, in terms of both size and statistical significance, to the 2SLS results in Table 2.

[31] $F[58, 467] = 66.40$ and $F[58, 467] = 23.87$ tests reject the null hypothesis that the dummies should be omitted in the OLS model with fixed effects for white and black teens, respectively. Results are similar for the 2SLS and GMM models.

[32] The primary difference is that the significance level drops for male wages in the results displayed in Table 3. The coefficient on male wages becomes insignificant at even a 10% level in columns 2 and 4.

(ii) Whether the proposed instruments are correlated with the endogenous variable. In this case, the question is whether the extra instruments are correlated with the benefit level.[33]

(iii) Whether instrumental variable estimation is needed at all (i.e., whether benefits are endogenous). If the benefit level is exogenous, then instrumental variables estimation will be less efficient than OLS, even when the instruments are valid.

For both black and white teens, the public choice variables perform satisfactorily on all three counts. Hansen (1982) tests of the overidentifying restrictions fail to reject the null hypothesis that the instruments are uncorrelated with the error term for both white and black teens (the χ^2 [2] statistics are 0.62 and 0.94, respectively). The partial F-statistic [3, 465] on the three public choice variables from the first stage regression is 6.95. This is not as large as one might wish but does indicate that the instruments are highly correlated with benefits. To test whether the benefit variable is endogenous, Wu–Hausman tests (Wu 1973; Hausman 1978) are performed. The tests compare the OLS and 2SLS coefficients on the (possibly) endogenous benefit variable. Under the null hypothesis that the AFDC benefit level is uncorrelated with the error term, OLS is efficient and consistent while 2SLS is merely consistent. Under the alternative, that the benefit level is correlated with the error term, OLS is inconsistent while 2SLS remains consistent. The χ^2 (1) statistics are 11.20 for white teens and 16.06 for black teens, rejecting the null hypotheses and indicating that 2SLS is appropriate.[34]

These results confirm that the public choice instruments are reasonable and that OLS fixed effects estimation is not appropriate. Although the tests of overidentifying restrictions fail to reject the null hypothesis that the instruments are valid, as we noted earlier, there might be reasons to suspect that per capita income remains correlated with illegitimacy rates, even after controlling for wages at the bottom of the incomes distribution. However, including per capita income in the base regression and dropping the Medicaid matching rate as an instrument does not affect results for either black or white teens. Results from this regression are shown in columns 7 and 8 of Table 2.[35] For white teens, the coefficients on government benefits and female wages remain statistically significant with the theoretically expected signs.[36] For black teens, the coefficient on government benefits remains statistically significant and the coefficients on the wage variables remain statistically insignificant.[37]

As an additional check for robustness, we also try a different set of instruments similar to variables used to explain AFDC benefits in Ribar and Wilhelm (1996).[38] The instruments are the difference between the percentage of the state's population and AFDC caseload that is black

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[33] The issue is whether the instruments not included in the second-stage regression are correlated with the endogenous variable in the first-stage regression (see Staiger and Stock [1997] or Pagan and Jung [1993]).

[34] T_3 from Bowden and Turkington (1984).

[35] GMM estimates are similar for this regression also.

[36] However, this is not because the percentage of the population over 65 is driving the results when all three variables are used as instruments. Results for both white and black teens are similar to those in columns 5 and 6, in terms of both size and statistical significance, when the percentage of the population over 65 is dropped as an instrument (using per capita income and the matching rate as the instruments) and when per capita income is included in the regression with the matching rate serving as the sole instrument.

[37] The coefficient on government benefits is statistically insignificant when states with few black teenagers are omitted from the regression for black teens (when the percentage of the population over 65 is the sole instrument).

[38] In Ribar and Wilhelm (1996), both the percentage of the AFDC caseload that is black and administrative costs were statistically significantly correlated with benefit levels in some specifications when regional, rather than state, dummies were included in the regression. They found that high program overhead and largely black caseloads were significantly negatively correlated with benefit levels.

Table 3. Further Results from Empirical Model (Using GMM with State and Time Dummies)

	(1)	(2)	(3)	(4)
Dependent Variable	Illegitimacy rate for white teens	Illegitimacy rate for black teens	Illegitimacy rate for white teens	Illegitimacy rate for black teens
Regression			Omitting male wages	Omitting male wages
Instruments	Public choice instruments[a]	Public choice instruments[a]	Public choice instruments[a]	Public choice instruments[a]
AFDC, Food Stamps, and Medicaid	0.0457[c] (2.31)	0.3510[c] (2.37)	0.0383[c] (2.59)	0.2680[c] (2.53)
Difference in Government Benefits				
Female Wage	−0.0275[c] (−2.43)	−0.0493 (−0.58)	−0.0267[c] (−2.41)	−0.0788 (−1.03)
Male Wage	−0.0038 (−0.69)	−0.0699[b] (−1.66)		
Abortion	−3.0586 (−1.30)	−0.2601 (−0.02)	−2.9251 (−1.31)	−0.2831 (−0.02)
Unemployment Rate	−0.2123[c] (−2.22)	−2.2101[c] (−2.62)	−0.1872[c] (−2.33)	−1.7866[c] (−2.61)
Number of Observations	535	535	535	535
Overidentifying Restrictions χ^2 (2)	0.5956	0.284	1.3830	0.442

t-statistic in parentheses.
[a] Public choice instruments are per capita income, percentage of the population over 65, and the Medicaid matching rate.
[b] Significant at the 10% level.
[c] Significant at the 5% level.

and administrative costs per AFDC family.[39] The difference between the racial composition of the recipient population and the state population might affect benefit levels if voters are less sympathetic toward recipients of different races than their own. Administrative costs might affect benefit levels since they affect program costs and might affect the public's perception about program efficiency. Since these variables probably do not affect the teens' decisions directly, they might be plausible instruments. Results using these instruments are shown in columns 9 and 10 of Table 2. For white teens, the coefficient on government benefits remains positive and is significant at a 5.2% level, while the coefficient on female wages remains negative and statistically significant. The point estimates are larger than in columns 5 and 7. The results for black teens are less encouraging—the coefficient on government benefits becomes statistically insignificant but remains positive. Results for both white and black teens are similar when per capita income is included directly in the regression (using the AFDC program variables as instruments).[40] Tests of overidentifying assumptions fail to reject these instruments at conven-

[39] Ribar and Wilhelm (1996) use administrative expenditures as share of total program expenditures. However, this variable was not readily available for the entire period in our study (1980–1990).

[40] When the percentage of the population over 65 and the AFDC program variables are used as instruments, the results are similar to those reported in columns 7 and 8 (when only the percentage of the population that is over 65 is used). For whites, the point estimate of the coefficient on government benefits is 0.08 and is significant at a 1% level and the coefficient on female wages is −0.05 and is also significant at a 1% level. For blacks, the coefficient on government

Table 3. Extended

(5)	(6)	(7)	(8)	(9)	(10)
Illegitimacy rate for white teens Abortion endogenous Public choice instruments[a]	Illegitimacy rate for black teens Abortion endogenous Public choice instruments[a]	Illegitimacy rate for white teens Abortion dropped Public choice instruments[a]	Illegitimacy rate for black teens Abortion dropped Public choice instruments[a]	Illegitimacy rate for white teens Difference included Public choice instruments[a]	Illegitimacy rate for black teens Difference included Public choice instruments[a]
0.0375[c]	0.2665[c]	0.0381[c]	0.2680[c]	0.0592[b]	0.2724[b]
(2.37)	(2.41)	(2.60)	(2.52)	(1.75)	(1.92)
				−0.0515	−0.0137
				(−0.70)	(−0.05)
−0.0192	−0.0344	−0.0259[c]	−0.0787	−0.0303[c]	−0.0782
(−1.12)	(−0.31)	(−2.36)	(−1.03)	(−2.31)	(−1.02)
23.3956	168.9752			−4.1375	−0.5373
(0.63)	(0.62)			(−1.45)	(−0.04)
−0.2387[c]	−2.1176[c]	−0.1922[c]	−1.7882[c]	−0.2249[c]	−1.7933[c]
(−2.14)	(−2.31)	(−2.42)	(−2.61)	(−2.15)	(−2.52)
535	535	535	535	535	535
0.4232	0.0000	0.9270	0.439	0.3335	0.442

tional levels (see Table 2). However, the AFDC program variables are less attractive in at least one way—they are less highly correlated with benefit levels than the public choice variables. The partial F-statistic in the first stage regression is 2.95, with a significance level of 9.8%.

Several variables have been consistently insignificant throughout the preliminary analysis. In particular, the female unemployment rate, the infant mortality rate, the share of the population living in metropolitan areas, and the incarceration rate are both singly and jointly insignificant for white teens (with an $F[4, 467]$ statistic of 0.6184). For black teens, two of the variables are significant at a 10% level, although they are jointly insignificant ($F[4, 467]$ statistic of 1.91). Keeping female unemployment and infant mortality in regressions for black teens similar to those in Table 3 does not affect the magnitude or the statistical significance of the coefficient on the benefit variable. However, the male wage variable's significance level drops below 10% when these variables are included. Excluding the four variables from the regression for white teens results in the coefficient on total unemployment becoming significant and negative. The negative sign (implying that high unemployment is correlated with low rates of illegitimacy) is hard to interpret in the context of the theoretical model. Reverse causality seems unlikely since

benefits is 0.36 and is significant at a 3% level. Once again, results for both black and white teens are similar when per capita income is included in the regression. Finally, when all the public choice and AFDC program variables are used simultaneously as instruments, the results are similar to those reported in columns 5 and 6 in terms of both size and significance. The only difference is that the coefficient on government benefits in the regression for black teens is smaller (0.22). In all regressions using the program variables as instruments (including regressions with additional public choice instruments), we continue to fail to reject the null hypothesis that the overidentifying restrictions are valid, at at least a 10% level.

most of the teens in the sample would not be in the work force even if not pregnant. Further, female teenagers do not make up a large portion of the work force and so small increases in illegitimacy rates would not affect the unemployment rate.

Estimating the reduced regression in a GMM framework, allowing for heteroscedasticity of unknown type, gives similar results to the larger regressions in Table 2 (Hansen 1982).[41] For white teens, female wages and government benefits are significantly correlated with teen illegitimacy rates in the expected directions, and the coefficient on male wages remains insignificant (see Table 3, column 1). For black teens, only the coefficient on benefit levels is consistently statistically significant with the theoretically expected positive sign. The coefficient on the female wage variable remains statistically insignificant throughout the entire analysis, and, as noted above, the male wage variable is not robustly correlated with the illegitimacy rate at conventional levels (column 2).[42]

In 1990, the simple correlation between male and female wages is 0.80, and across the whole sample, the simple correlation between the two is 0.74—high enough to make multicollinearity a concern. However, dropping male wages from the regression has little effect on results for either white or black teens (see columns 3 and 4). For black teens, it increases the significance of female wages, although the coefficient remains insignificant at conventional levels.[43]

One question we have not discussed is the appropriate way of dealing with the proxy for abortion services.[44] The abortion proxy is the share of counties in the state that have an abortion provider. This variable is quite possibly endogenous because areas with low rates of teen pregnancy might have little demand for providers.[45] When the abortion variable is dropped, the results are similar to earlier results for both black and white teens (Table 3, columns 7 and 8). This regression is repeated treating the abortion variable as endogenous (Table 3, columns 5 and 6). This has little effect on the coefficient on government benefits for either white or black teens. The coefficient on female wages becomes insignificant for white teens. However, the partial F-statistic on the extra instruments in the first-stage abortion variable regression indicates the public choice instruments have little explanatory power for this variable.[46]

Another question is what effects additional increments in government benefits for additional children have on illegitimacy among teens. The theoretical section indicates that additional increments are more likely to encourage second or later births but that increases in increments may also increase the chance of a first birth. This is because increases in increments affect the budget constraint when the woman chooses welfare. However, the level of benefits for a family of two may be more relevant if teens are not forward looking. Table 3, column 9 shows that the results for white teens are similar to earlier results except that the significance level of the base payment drops to a 10% level. However, the sign of the coefficient on incremental benefits is counterintuitive. The result for black teens is also similar (see Table 3). These results might

[41] Two-stage least squares results are similar for the reduced regression.

[42] These results are also robust to dropping per capita income and the Medicaid matching rate as instruments and including per capita income in the regression.

[43] Excluding female wages from the regression for black teens does not have a large effect on the coefficient for male wages. The coefficient remains significant at a 10% level when the female unemployment and infant mortality rates are excluded from the regression and remain insignificant when they are included in the regression.

[44] Jackson and Klerman (1994) and Kane and Staiger (1996) address the issue of the effect of abortion on teen childbearing.

[45] Since many older women also have abortions, this does not necessarily follow.

[46] The $F[3, 474]$ statistic is 0.794.

Table 4. Lagged Dependent Illegitimacy Rates Included for White and Black Teens[a]

Dependent Variable[b]	First difference of illegitimacy rate for white teens	First difference of illegitimacy rate for white teens	First difference of illegitimacy rate for black teens	First difference of illegitimacy rate for black teens
Lagged Illegitimacy Rate	−0.0538		0.0329	
	(−0.10)		(0.43)	
Government Benefits	0.0238[c]	0.0228[d]	0.0251	0.0263
(family of 2)	(1.82)	(2.20)	(1.39)	(0.29)
Female Wage	−0.0108	−0.0123	0.0014	0.2050[d]
	(−0.79)	(−0.88)	(0.05)	(2.07)
Abortion	−3.9232	−4.2993	−16.721	64.730
	(−0.52)	(−0.56)	(−0.66)	(1.03)
Unemployment Rate	0.0744	0.0491	0.0254	0.3059
	(0.33)	(0.29)	(0.10)	(0.32)
Number of Observations	423	423	423	423

t-statistic in parentheses.
[a] All variables are first differences. Estimation includes time dummies.
[b] First difference fixed effects model (Equation 8) from Keane and Runkle (1992a).
[c] Significant at the 10% level.
[d] Significant at the 5% level.

be due to multicollinearity caused by the high correlation between base and additional payments. The partial correlation between the two variables is 0.83.

As a final exercise, we consider the effects of lagged illegitimacy rates.[47] It seems reasonable that changes in the prevalence of out-of-wedlock childbearing may effect attitudes and thus effect future rates. In a short panel, such as the one used here, it is not possible to test this hypothesis fully due to the limited number of time periods available. However, to test a limited form of this hypothesis, we include a single lag of the illegitimacy rate. As noted in Keane and Runkle (1992a), first differencing, rather than taking the fixed effects transformation, eliminates the individual state effects and allows instruments to be predetermined rather than strictly exogenous.[48] In particular, Keane and Runkle (1992a) propose a GLS estimator that eliminates possible serial correlation in the error term u_{it} with a forward-filtering transformation and only requires the instruments to be predetermined.[49] For white teens, the results appear consistent with the earlier results, government benefits remain negative and significant, and the coefficient on female wages remains negative but is insignificant (see Table 4). Since the coefficient on lagged illegitimacy rates is small and statistically insignificant, column 2 drops the insignificant illegitimacy rate. For black teens, none of the coefficients on the independent variables are

[47] Winegarden and Bracy (1997) explore this issue using aggregate U.S. data from 1973–1992. They find that lagged illegitimacy has a statistically significant effect on illegitimacy. They suggest that this variable is a control for cultural change.

[48] Strict exogeneity requires that $E(u_{it}z_{is}) = 0$ for all s and t, whereas predetermined only requires $E(u_{it}z_{is}) = 0$ for $s \leq t$. First differencing makes i.i.d. errors follow an MA(1) process.

[49] This requires instruments from period $t - 1$ or earlier. We use lagged first differences of the exogenous variables and the public choice instruments and allow for the errors to follow a more general MA process (as described in Keane and Runkle 1992a). In general, the estimation method Keane and Runkle (1992a) propose is not as efficient as a GMM estimator proposed by Arellano and Bond (1991). However, the Arellano and Bond (1991) GMM estimator uses all lags of all predetermined variables (and all leads and lags of strictly exogenous variables) as instruments. This means there are literally hundreds of moment conditions. In this application, where there are only 423 observations, this estimation method is not practical (see Chamberlain 1992; Keane and Runkle 1992b).

Table 5. Estimated Elasticities at Means of Variables

	Table 2, Column 3 (OLS–FE)	Table 3, Column 1 (GMM)	Table 3, Column 3 (GMM, Omitting Male Wages)
White Teens			
Government benefits	0.18[a]	1.53[b]	1.28[b]
Female wages	−0.16	−0.38[b]	−0.37[b]
Male wages	0.05	−0.07	
Black Teens			
Government benefits	−0.15	2.72[b]	2.08[b]
Female wages	0.36	−0.16	−0.25
Male wages	0.01	−0.31[a]	

[a] Significant at the 10% level.
[b] Significant at the 5% level.

statistically significant when lagged illegitimacy rates are included in the regression. When the lagged illegitimacy rate is dropped from the regression, the coefficient on female wages becomes statistically significant with a theoretically inconsistent positive sign (see Table 4).[50] Overall, these results give no support to the hypothesis that increased illegitimacy among teens leads to changes in attitudes that then lead to further increases in illegitimacy. However, this might be due to the very limited form of hypothesis that is being tested.

Summary of Estimation Results and Implications

The results of the econometric investigation are summarized below. Table 5 shows elasticities, computed at the means of all variables, for results from Table 2, columns 3 and 4; Table 3, columns 1 and 2; and Table 3, columns 3 and 4).

The basic results from the empirical estimation are as follows.

(i) Including fixed effects in the regression, but otherwise treating AFDC as exogenous, leads to parameter estimates for white teens that are consistent with theory but imply very modest elasticities of illegitimacy with respect to the variables of interest (0.18 for benefits and −0.16 for female wages). For black teens, the parameter estimates are inconsistent with theory, although the government benefit variable's parameter is statistically insignificant at conventional levels. These results—small but positive elasticities for white teens and inconsistent and statistically insignificant results for black teens—are similar to results in other studies that have used state-level data and fixed effects estimation to study the effect of welfare payments on illegitimacy among teenagers (e.g., Jackson and Klerman 1994). However, Wu–Hausman tests strongly reject the null hypothesis that AFDC benefits are uncorrelated with the error term from the illegitimacy regression for both black and white teens. This strongly indicates that fixed effects estimation does not resolve all endogeneity concerns in these regressions.

[50] Results are similar, but not identical, when this model is estimated in a GMM framework using first differences and lags of first differences of the exogenous variables and public choice instruments as instruments (and allowing the error term to follow a general MA process). For white teens, benefits remain significant, female wages become significant, and the coefficient on lagged illegitimacy rates remains insignificant and positive. For black teens, the lagged illegitimacy rate has a counterintuitive negative sign (indicating that high illegitimacy rates last year are correlated with low illegitimacy rates this year) and all other independent variables are insignificant.

(ii) The instruments used for the benefit variable are similar to those suggested in the public choice literature on the determinants of state welfare policy—per capita income, the Medicaid matching rate, and the percentage of the population that is over 65. Tests of overidentifying restrictions fail to reject these variables at conventional levels. These instruments are highly correlated with the benefit levels. One concern, despite the test of overidentifying restrictions, is that per capita income might not be an appropriate instrument. However, as shown in the results section, including per capita income in the regression and dropping the Medicaid matching rate as an instrument (since this is a function of per capita income) does not affect the results for either black or white teens.

(iii) For both white and black teens, the combined value of AFDC, food stamps, and Medicaid benefits for a family of two are positively correlated with illegitimacy rates. Results are similar when just AFDC and food stamp benefits are used.[51] The point estimates of elasticities (see Table 5) are 1.28 and 2.08 for white and black teens, respectively.[52] This is larger than point estimates in studies that have not taken the endogeniety of benefits into account. However, the 95% confidence intervals are large ([0.47, 2.10] and [0.73, 3.43], respectively). This result is highly robust for white teens but is slightly less robust for black teens. In particular, for black teens, the coefficient on benefits becomes statistically insignificant when lagged illegitimacy rates are included in the regression and when an alternate set of instruments is used. The incremental benefit for additional children is insignificant for both black and white teens, but this may be because it is highly correlated with the benefits for a family of two.

(iv) The measures of male and female wages, used in this paper, are the median weekly wages for persons working full-time, aged between 21 and 35, with a high school diploma or less. For white teens, female wages are robustly negatively correlated with illegitimacy rates. The point estimate of the elasticity is -0.37, with a confidence interval of $(-0.62, -0.12)$. Male wages are not statistically significantly correlated with illegitimacy rates. For black teens, male wages are statistically significant at a 10% level in some specifications, although this result is not highly robust. Female wages are consistently statistically insignificant throughout the analysis for black teens.

(v) There is no evidence that increases in illegitimacy lead to changes in attitudes that then lead to further increases in illegitimacy. However, we were only able to test a limited form of this hypothesis due to a lack of data.

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[51] The main differences are as follows.

(i) *For White Teens.* When per capita income is included in the regression, using the percentage over 65 as the sole instrument, the significance level on female wages tends to drop. In particular, in Table 2, column 8, the coefficient on female wages is only significant at a 10% level.

(ii) *For Black Teens.* The significance level of the coefficient on male wages tends to increase. In Table 2, column 6, it becomes significant at a 10% level and in Table 3, column 2, it becomes significant at a 5% level. In the smaller model, when per capita income is included in the regression, the coefficient on benefits drops to a significance level of 10%.

(iii) *For Both Black and White Teens.* In the regression where abortion is treated as endogenous, (Table 3, columns 5 and 6), the significance level of the coefficient on the benefit (AFDC and food stamps) variable drops to 10%. In the regression where the difference between AFDC and food stamps for a family of four and a family of two is included (Table 3, columns 9 and 10), the significance level of the coefficient on the benefit variable (AFDC and food stamps) increases to a 5% level.

[52] The point estimate of the elasticity for the sample of states that excludes states with fewer than 2000 black teens is smaller (1.10). This falls in the 95% confidence interval for the full sample of states for black teens. As noted earlier, a Chow test fails to reject the null hypothesis that the coefficients are the same for states with many and few black teens.

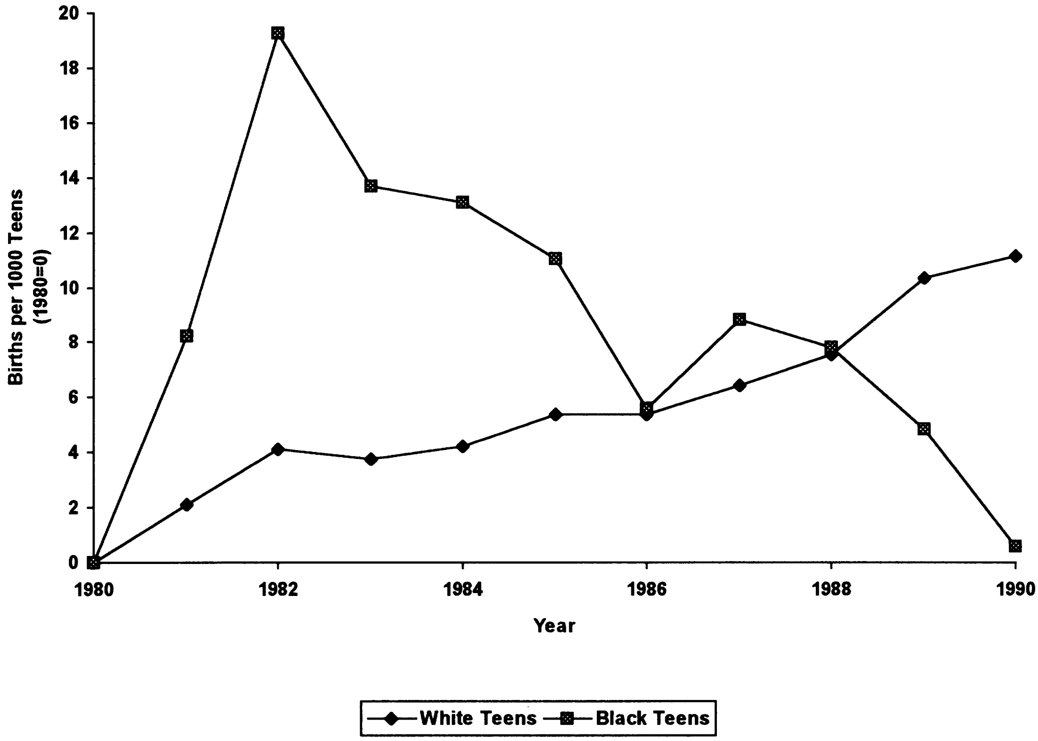

Figure 2. Time Dummies

(vi) Other control variables (including female unemployment rates, share of counties in the state with an abortion provider, infant mortality rates, incarceration rates, and share of the population living in metropolitan areas) are generally statistically insignificant throughout the analysis. The unemployment rate has a counterintuitive, negative correlation with illegitimacy rates, implying high unemployment is correlated with low illegitimacy rates. Reverse causation seems unlikely because female teenagers only make up a small portion of the workforce.

(vii) The independent variables do not fully explain the trends in illegitimacy over the period studied. For white teens, the coefficients on time dummies generally trend upward over the period, although they were fairly flat between 1982 and 1986. In contrast, the coefficients on the time dummies for black teens are largest between 1982 and 1984 and then generally trend downward (see Figure 2).

Overall, the results for both white and black teens are remarkably consistent with theory and stress the importance of economic incentives on the choice between work and welfare. A 1% increase in welfare benefits appears to increase illegitimacy among both white and black teens by more than 1%. A 1% increase in female wages appears to have a more modest effect of about a 0.4% decrease in illegitimacy for white teens but does not appear to affect illegitimacy rates for black teens.

The results also stress that fixed effects estimation alone does not appear to control for the endogeneity of benefits in aggregate data. Fixed effects estimation treating government benefits as exogenous yields much smaller estimates of elasticities for white teens and results that are inconsistent with theory for black teens. Several points should be made with regard to the results from the empirical estimation. The first is that the results are only for illegitimacy among teens and should not be extrapolated to other groups or to other questions. An increase in illegitimacy among teens may only have a modest effect on female headship, even in the long run. Furthermore, it is possible that, even though government benefits might affect illegitimacy among teens, the effect on other family decisions, such as out-of-wedlock childbearing among older women or divorce among married women, might be far smaller. Second, the results do not imply that benefits necessarily have as large an effect on teen childbearing as they do on illegitimacy. The positive correlation could be due to benefits discouraging marriage among pregnant teens rather than encouraging births.

6. Conclusions

The observation that the U.S. welfare system might encourage fertility among poor single women by making children income-producing assets as well as consumption goods is not novel (see, e.g., Becker 1991). Nevertheless, past empirical work, which has used differences in benefit levels over time and across states to test this hypothesis, has found that the effect is weak, inconsistent, and often statistically insignificant (see Moffitt 1992 or Murray 1993). However, benefit levels in state welfare programs are not the result of a natural experiment—politicians and voters choose the benefits that their state offers. If voters' perceptions about welfare dependency or illegitimacy affect their preferred benefit levels, then benefits and out-of-wedlock births will be codetermined. If so, coefficients from fixed effects estimation (which is a form of ordinary least squares) will be biased and inconsistent. Hypothesis tests confirm that benefit levels are endogenous. Once we control for this, we find large and statistically significant results for both black and white teens. These results are robust to several different instrument choices. In summary, we find the following.

(i) As noted by Ellwood and Bane (1985), many omitted, and potentially unmeasurable, state characteristics might affect both benefit levels and illegitimacy. If these characteristics are (roughly) constant for each state across the entire period, then including state dummy variables might allow unbiased estimation of coefficients in the illegitimacy regression. However, hypothesis tests suggest that including state (and time) fixed effects does not adequately control for endogeneity. For white teens, the coefficients from standard fixed effects estimation are consistent with theory and are statistically significant but small (elasticities of about 0.2). For black teens, results from standard fixed effects estimation are inconsistent with theory and are statistically insignificant. Once endogeneity is controlled for, welfare benefits are strongly and robustly related to teen illegitimacy for both white and black teens, with elasticities of around 1.3 and 2.1, respectively.

(ii) Female wages, for women aged between 21 and 35, working full-time, with a high school education or less, are robustly correlated with illegitimacy rates among white teens, with an elasticity of around -0.4. Female wages are not significantly correlated with illegitimacy rates for black teens. Male wages (for the same group) do not appear correlated with illegitimacy rates among white teens. For black teens, male wages are statistically significant (with the expected negative sign) in some specifications, but this result is not highly robust.

A number of theoretical and empirical questions remain open and deserve further investigation. Although the coefficients on the wages and benefits have the correct theoretical signs, the variables do not entirely explain the rapid growth in illegitimacy over the past 10 years. For white teens, time dummies are highly significant and increasing over this time period.

Explaining this growth in a more satisfactory, and testable, manner than changes in attitudes would seem an important goal for future research.[53]

Because the dependent variable is the illegitimacy rate, not the birth rate, it is important to note that a positive coefficient on government benefits (and a negative coefficient on wages) does not necessarily imply benefits encourage (or wages discourage) teens to get pregnant. The signs could be primarily the result of fewer pregnant teens getting married rather than more teens bearing children. Finally, extending the research back to the 1970s might be useful. However, this may be difficult because of changes in the food stamp program and abortion access during this time period.[54]

Appendix A: Data Definitions and Sources

Means and variances of all variables are presented in Table A1. The sources of the data are listed below.

Incarceration Rates. Number of sentenced prisoners in the state per 100,000 resident population. United States Department of Justice, Bureau of Justice Statistics, *Sourcebook of Criminal Justice Statistics,* 1991.

Abortion. Percent of counties in a state with an abortion provider of five or more abortions. Alan Guttmacher Institute, *Abortion Factbook: 1992 Edition.* Data for 1989 and 1990 are the number of providers in 1988.

Wages. Data for weekly wages for women and men are computed from data from the National Bureau of Economic Research's CPS Labor Extracts (2nd edition) on CD-ROM. They are the (weighted) median of weekly wages for persons with a high school education or less, working full-time (over 35 hours a week), aged between 21 and 35 of the relevant sex.

Prices are inflated to 1991 prices by using the June consumer price index for urban consumers.

Share of Population Living in Metropolitan Areas. The percent of the population living in MSA's in that year. Data for 1981–1982 are interpolated. *Statistical Abstract of the United States,* various years.

Infant Mortality. Deaths per 1000 infants in the first year of life. *Vital Statistics of the United States,* various years.

Unemployment Rates (Both Women's Unemployment and Total Unemployment). Unemployment rate. 1980, Bureau of Labor Statistics bulletin "Geographic Profile of Employment and Unemployment." 1981, Bureau of Labor Statistics Bulletin 2175, "Handbook of Labor Statistics, 1983." 1982–1989, Bureau of Labor Statistics Bulletin 2340, "Handbook of Labor Statistics, 1989." Data for 1990 are from the "Geographic Profile of Employment and Unemployment," 1990.

Illegitimacy Rates. The number of illegitimate births per 1000 single females aged 15–19.

The number of illegitimate births is from *Vital Statistics of the United States,* various years. The estimate of the number of females is calculated as follows. The number of females in each state in the single year groups for 15- through 19–year olds is from Bureau of Census, Current Population Division estimates. (These figures were corrected by the Bureau of Census to be consistent with the 1980 and 1990 census population counts.) To get the number of females by race, the percentages of females who are white and black in each year group are interpolated between census years. These percentages are multiplied by the total number of females in that year group, and the resulting number by year group are summed to get a total for females aged 15–19.

AFDC, Food Stamps, and Medicaid Benefits. The total monthly payment of AFDC and food stamps to a family with two members and no other income and average Medicaid expenditures for one adult and one child in the AFDC program. The incremental payment is the difference between the monthly payment for a family of four and a family of two. It is assumed the additional members are children. The food stamp payment is calculated assuming that the maximum shelter deduction and the standard deduction were used when calculating payments. AFDC and food stamp data were provided by the Congressional Research Service. Medicaid data were provided by the Division of Medicaid Statistics at the Health Care Financing Administration.

Per Capita Income. Per capita income. *Statistical Abstract of the United States.*

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[53] See, for example, Nechyba (1997) for a theoretical model of illegitimacy and the AFDC program with changes in attitudes.

[54] Prior to 1977, food stamp recipients were required to buy food stamps at a discount of their face value; this was abolished because of concerns that individuals were not always able to afford their allotments (Ohls and Beebout 1993). Roe v. Wade in 1973 and the Hyde amendment in 1976, which ended federal Medicaid funding of abortion, may also be important.

Table A1. Means of Variables

	Mean	Standard Deviation
Dependent Variables		
Illegitimate births per 1000 white teens	19.64	6.95
Illegitimate births per 1000 black teens	85.02	28.71
Independent Variables		
AFDC, food stamps, and Medicaid (family of 2) ($)	659.46	136.14
AFDC, food stamps, and Medicaid (difference) ($)	368.93	53.41
Weekly female wage ($)	273.59	33.08
Weekly male wage ($)	377.25	55.99
Percent of counties with an abortion provider	0.278	0.27
Unemployment rate	6.921	2.38
Incarceration rate	175.9	85.93
Infant mortality rate	10.54	1.75
Percent of population in metropolitan areas	62.77	22.37
Unemployment rate for women	6.98	2.30
Instruments		
Medicaid matching rate	0.560	0.09
Percent of population over 65	0.117	0.02
Monthly per capita income ($)	1338	232.79

Medicaid Matching Rate. Federal matching rate for the medicaid program. *Characteristics of State Plans for AFDC,* various years.

Percent of Population over 65. The percent of the population aged over 65. Data were provided by the Bureau of Census, Current Population Estimates Division.

Percentage of the AFDC Caseload That Is Black. The percent of the AFDC caseload that is headed by an individual who is Black. Data are from Committee on Ways and Means "Overview of Entitlement Programs" (various years). Data for 1981–1982 and 1983–1984 are interpolated. Data for 1980 are missing for several states and so these observations are dropped when this variable is used as an instrument.[55]

Administrative Costs per AFDC Case. State administrative expenditures per AFDC case. Data are from Committee on Ways and Means "Overview of Entitlement Programs" (various years).

Appendix B: Sample Sizes in Individual Studies

Individual studies of the effects of AFDC on teen illegitimacy tend to use data from two sources. The first is the National Longitudinal Survey of Youth (NLSY) and the second is the Panel Study of Income Dynamics (PSID). As Plotnick (1990, p. 738) notes, "differences in [out-of-wedlock childbearing] behavior can not be adequately captured by inclusion of race/ethnicity dummies since, as several studies have shown, the effects of explanatory variables tend to differ among the groups." As a result, samples are often separated for black and white teens. Once divided, the typical sample size in published studies is often between 500 and 1000 individual teens.[56]

However, out-of-wedlock births to white teens are relatively rare events, and although more common among black

[55] Results in columns 9 and 10 in Table 4 are similar in terms of both size and significance when interpolated points are dropped from the regression.

[56] See Duncan and Hoffman (1990), Plotnick (1990), and Lundberg and Plotnick (1994). Using the Current Population Survey or the Public Use Micro Sample (PUMS) from the Census Bureau would increase sample size. However, this would mean much panel information would be lost. The data loss would be most severe for teens who give birth and then move out of their parent's household since family background data, which in general provides the best available proxies for lifetime wages, will be lost. Omitting teens who moved out of their parents' households might exclude those most likely to have been influenced by the availability of government benefits.

Table B1. Sample Sizes, Number of Births, and Number of States in NLSY Samples

	Number of Observations[a]	Number of Births	Number of States
White Teens			
Family income included	561	44	22
Base regression	688	56	23
Black Teens			
Family income included	281	94	23
Base regression	322	102	23

For women who gave birth out-of-wedlock as a teenager (before 20th birthday).
[a] Includes teens in states where no births are observed.

teens, births are still only observed in a few states since the black population is more concentrated and sample sizes tend to be smaller.[57] This complicates estimation since, when there are no out-of-wedlock births in a state, the state dummy can perfectly "predict" nonbirths by becoming arbitrarily, negatively large.[58] This reduces variation in the AFDC variable and decreases actual sample size since only teens from states where births are observed can be included in the regression. Table B1 shows births to white and black teens for a sample from the NLSY. It shows total observations (including observations in nonbirth states), total out-of-wedlock births, and the number of states where teens gave birth. This last column gives information on how many states would be excluded by including fixed effects in the regression. The teens in this sample are all the teens who turned 15 or 16 in 1979. This is essentially the whole sample that is observed during their entire (fertile) teenage years and is the sample used in Plotnick (1990).[59]

To take account of missing independent variables, we only count observations for which data on family income in 1978, mother's education level, and family structure at age 16 are available. As noted, 44 white teens from only 22 states gave birth before their 20th birthday in this sample (out of 561 teens).[60] Dropping family income (in general this variable is often missing even when interviews took place) increases the sample to 688 cases with 56 total births. However, even in this larger sample, the births occur in only 23 states.

Including older teens, for example teens who turned 17 in 1979, increases sample size.[61] However, this means that many control variables will not be observed during the woman's teen years. Further, even in these larger samples, the number of states where births occur remains small. For example, when no other regressors are included, although sample size for white teens is increased to over 1000 individuals and 85 births, the births occur in only 25 states.

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[57] For black teens, there are also several states where there are no women who did not give birth as teenagers.

[58] For studies of the effect of total out-of-wedlock fertility (e.g., Rosenzweig 1995), it would be possible to follow the teens for longer periods (i.e., to age 23 or age 26). This would increase the number of births (both by increasing the length of time the women are at-risk of having an out-of-wedlock birth and because out-of-wedlock births are more common among women in their early 20s than among women in their teens). However, as noted earlier, this is a slightly different policy question given that older women are less likely to become dependent on welfare.

[59] Plotnick (1990) has sample sizes of 488 individuals for white teens and 230 individuals for black teens. Plotnick (1990) includes different explanatory variables in the regression, which might result in more missing values.

[60] Sample sizes include teens who would be dropped from the regression if state dummies were included in the regression.

[61] For example, Lundberg and Plotnick (1994) include teens who turned 17 in 1979.

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