Multiple-Father Fertility and Welfare

Evenhouse, Eirik and Reilly, Siobhan

Mills College

23 October 2010

Multiple-Father Families and Welfare

Eirik Evenhouse

Department of Economics

Mills College

Home phone: 510-654-5653

Cell phone: 510-301-2903

Eirik.evenhouse@gmail.com, eevenhou@mills.edu

Siobhán Reilly

Department of Economics

Mills College

Home phone: 510-654-5653

Cell phone: 510-301-2903

Siobhan.reilly@gmail.com, sreilly@mills.edu

Authors' acknowledgements

For comments and advice, the authors thank Ronald Mincy, Robert Moffitt, Steve Raphael, two anonymous reviewers, and participants at the 2006 meetings of the Population Association of America. The authors acknowledge the financial support of the National Institute for Child Health and Human Development, under grant R03HD39933.

ABSTRACT *(149 words)*

In the United States, multi-partnered fertility (MPF) has become commonplace. This study provides the first nationally representative measures of women's MPF, across multiple years, using the U.S. Census Bureau's Surveys of Income and Program Participation. Because welfare rules contain strong incentives for MPF, and because MPF is especially common among welfare recipients, we also examine the relationship between welfare and MPF. Focusing on the pre-TANF period 1985 to 1996, when welfare rules were more comparable across states and the absence of time limits made the incentives for MPF larger, we find little behavioral response. Among low-income mothers, MPF does not appear to be driven by program design. Because the incentives were relatively large and reached well up the income distribution, our findings amplify those of earlier studies that show little demographic response to antipoverty programs, and invite reconsideration of how much these incentives should constrain transfer programs that target children.

KEYWORDS: Multipartnered fertility, multiple-father family, SIPP, welfare, AFDC

I. Background

Multi-partnered fertility has become commonplace, especially among the poor and other vulnerable subgroups, yet two of its key aspects remain unexplored. It has not been measured among women in a large, nationally representative data set over multiple years, and few studies specifically target its causes. Meanwhile, there are reasons to suspect it is an important phenomenon, worthy of further investigation. Multi-partnered fertility creates practical challenges for families and knotty problems for policymakers, and its concentration among at-risk groups magnifies its importance. Although the causal relationships (if any) are not yet clear, it is correlated with a host of negative indicators of child and family wellbeing. However, its prevalence suggests it may also hold advantages (for example, expanding or diversifying sources of household income), but these, too, have had scant formal exploration.

One advantage large enough to be a credible incentive is an income premium created by the U.S. welfare system in favor of certain forms of multi-partnered fertility. This study exploits variation in welfare benefits -- over time and across states -- to examine the link between welfare and the prevalence of multi-partnered fertility.

Klerman (2007) argues that, for multi-partnered fertility to warrant the attention of policymakers, "it must be reasonably common and have important consequences." Certainly it is "reasonably common." In the 2002 National Survey of Family Growth (NSFG), 17 percent of all fathers aged 15-44 report having children by more than one woman (Guzzo and Furstenberg 2007a). In 2001 data from the National Longitudinal Survey of Adolescent Health (Add Health), 14 percent of mothers aged 19-25 whose first birth was nonmarital had had a subsequent birth with another partner (Guzzo and Furstenberg 2007b). In the Fragile Families and Child Wellbeing Study, 23 percent of mothers have children by more than one man (Carlson and Furstenberg 2006). Among unmarried couples in that survey, rates are even higher: in 22 percent of cases the father has children from a prior relationship, in 17 percent of cases the mother does, and in 20 percent both father and mother do (Roberts 2008). Among men aged 35-44 with incomes below 150 percent of the poverty line, the rate rises to 27 percent for whites and 37 percent for blacks, with 16 percent of black men in that category having children by three or more partners (Guzzo and Furstenberg 2007a).

Multi-partnered fertility may, too, to have "important consequences." It is, at any rate, negatively correlated with a number of factors believed to favor child and family wellbeing. For instance, the presence of children from prior relationships reduces the probability of marriage, for mothers as well as fathers. In a sample of low-income women in inner-city Philadelphia who had just given birth, "men who had children with multiple partners were significantly less likely to cohabit with or to be married to the mother of the focal child, net of demographic and socioeconomic characteristics" (Margolis and Mykyta 2008, p. 15). This finding is echoed in analyses of the Fragile Families survey (e.g., Mincy and Huang 2001, Carlson and Furstenberg 2006). Harknett and Knab (2007) find multi-partnered fertility negatively associated with mothers' perceptions of the availability of financial, housing, and child care support, and use longitudinal data to infer a bi-directional relationship between multi-partnered fertility and the mothers' support received from their social networks. Turney and Carlson (2010) suggest that parents' mental health and multi-partnered fertility may also be reciprocally related, and note that mothers and fathers who are depressed, and mothers who report heavy episodic drinking, are more likely to have a child by a new partner.

Multi-partnered fertility is also correlated with reduced investment in children in the form of father-child contact (e.g., Cooksey and Craig 1998; Manning and Smock 1999) as well as child support payments (e.g., Huang, Mincy and Garfinkel 2005; Manning, Stewart, and Smock 2003).

While researchers still know relatively little about the consequences of multi-partnered fertility, even less is known about the factors that give rise to it. The possible policy responses offered by Klerman (2007) – promoting marriage, supporting family planning, enforcing child support obligations, and improving the financial status of low-income fathers – hint at possible causes. The welfare system, however, seems an obvious candidate. Welfare rules create large financial incentives in favor of multi-partnered fertility. Moreover, its incidence among recipients is particularly high. Analyzing Wisconsin welfare data from 1997-1998, Meyer, Cancian, and Cook (2005) find that, among mothers of two or more children, 39 percent had children by more than one man. Even that figure is only a lower bound, as another 44 percent of the mothers in their sample had at least one child whose paternity had not been legally established. That prompts the question of whether welfare is a driver of multi-partnered fertility.

The behavioral consequences of welfare programs have long been of interest to social scientists and policymakers, and there is a large literature on the effects of welfare programs on family structure (see Moffitt 2003 for a recent survey). That literature usually finds little or no demographic response to the incentives created by welfare programs. However, the incentive in favor of multi-partnered fertility is larger, and reaches further up the income distribution, than those previously studied, making a response more likely. This study is a first step in that direction. If closer examination should establish a connection between welfare and multi-partnered fertility, one could infer that multiple-father fertility might be partly an economic coping strategy for low-income mothers. Conversely, finding that behavior to be unresponsive even to the large financial inducements present would constitute evidence that multiple-father fertility is a response primarily to other, non-financial factors. It would also call into question the extent to which policymakers should sacrifice other goals, such as income adequacy for children, to concern over demographic incentives.

The analysis focuses exclusively on the decade preceding the welfare reforms that took place after 1996, so that it is not confounded by any abrupt behavioral changes provoked by the reforms. Before Aid to Families with Dependent Children (AFDC) was replaced by Transitional Assistance to Needy Families (TANF), cross-state variation in benefit levels was the principal source of variation in the financial incentive for multi-partnered fertility; afterward, states' welfare programs became far more heterogeneous, making it far harder to place a dollar value on that incentive. Moreover, the introduction after 1996 of time limits and stricter work requirements shrank the incentive, further adding to the challenge of identifying behavioral responses.

In addition to considering the link between welfare and multi-partnered fertility, this study also offers a set of estimates of women's multi-partnered fertility spanning the 1985-1996 period, for a variety of subpopulations. These are the first time-series estimates for women to be derived from a large, nationally representative sample (the Survey of Income and Program Participation, or SIPP). Existing estimates draw most often on the Fragile Families survey, which is smaller, urban, and focused on nonmarital births. The estimates of Guzzo and Furstenberg (2007b) are based on Add Health, which includes only young women (aged 19 to 25 in 2001) and in which high school dropouts are underrepresented. Those of Meyer *et al.* (2005) are based on administrative data from one state's welfare program. The other estimates of which we are aware concern the multi-partnered fertility of fathers (e.g., Manning *et al.* 2003; Guzzo and Furstenberg 2007a).

In this study, we measure only women's multi-partnered fertility (henceforth, multiple-father fertility), for three reasons. First, we are interested in the contributory role of the welfare system and, while welfare's eligibility rules are gender-neutral, women predominate among adult recipients. Second, measuring men's multi-partnered fertility in nationally representative datasets is harder than measuring that of women. The handful of surveys containing the relationship detail needed to measure multi-partnered fertility indirectly are household-based surveys, and men usually live apart from their children from prior relationships. Only in relatively recent surveys have respondents been asked directly how many partners they have children by (the NSFG, for instance, did not introduce this question for male respondents until 2002). Finally, women's reports of how many children they have borne are considered more accurate than men's reports of the number of children they have fathered (Rendall *et al.*, 1999).

In the next section, we explain how welfare program rules favor the formation of multiple-father families, and in Section III we present empirical evidence concerning the hypothesized relationships between welfare and multiple-father fertility. Section IV concludes.

II. Welfare's Incentives for Multiple-Father Family Formation

There are two avenues by which welfare could encourage the formation of multiple-father families. First, welfare appears to increase the odds of family breakup, which, in turn, opens the way for the formation of a multiple-father family. Second, just as child-support laws may treat the multiple-father family more generously than they do a one-father family (Meyer *et al.* 2005), in some circumstances welfare rules also favor the multiple-father family.

Welfare's subsidy to family breakup is the result of conditioning eligibility on the absence of a biological parent. Until 1968, the presence of any adult man in the household barred a family from receiving AFDC benefits. In 1968, the Supreme Court ruled that cohabitation *per se* does not automatically disqualify a welfare recipient from benefits. In 1970, it ruled further that the cohabitor's income cannot be counted in determining the family's benefit unless caseworkers prove that the cohabitor supports the family. The standards of proof are difficult to meet (Moffitt, Reville, and Winkler 1994, 1998). Consequently, welfare bureaucracies have generally made it a very low priority to ascertain the contributions of a recipient's adult friends (Rolston 2000).

The Supreme Court rulings, together with the decline in social stigma attached to family breakup and unmarried cohabitation, meant that, throughout the 1985-1996 period examined here, the only man whose presence was an obstacle to welfare eligibility was the biological father of a woman's children.[1] A mother could get the companionship and the household scale economies that come with cohabiting and still receive welfare, a small but steady income that she neither had to work nor bargain for, and that made her more attractive in the eyes of potential partners. Not surprisingly, by 1990 a significant fraction of mothers on welfare were living with unrelated men, many of whom had substantial earnings (Moffitt *et al.* 1998, London 1998). There is also some evidence from that period that higher welfare benefits corresponded to a slightly higher likelihood that a mother was cohabiting with a man rather than married to him, and that an unmarried mother's male cohabitor was not related to her children (Evenhouse and Reilly 2005).

It should be emphasized that the AFDC-UP (Unemployed Parent) program did not eliminate welfare's implicit subsidy for replacing fathers with unrelated men. This is perhaps counter-intuitive, as AFDC-UP was created in 1961 to blunt the incentive created by AFDC for poor parents to separate. The AFDC-UP program meant merely that a subset of two-parent families—those meeting the income and work history criteria—were no longer *categorically* ineligible for welfare. However, in AFDC-UP as in the basic AFDC program, a biological father's income was taxed at an implicit marginal rate of virtually 100 percent, while the income of an unmarried, unrelated cohabitor was effectively untaxed. Thus, the expansion of AFDC-UP that occurred during the latter part of the 1980s—from 23 states in 1982 to all states by 1990—had little bearing on the incentive considered in this study.

It should also be emphasized that welfare's implicit subsidy for replacing fathers with unrelated men was large enough to be relevant for many households. Put differently, a large proportion of parents were paying a "tax" for living together (by forgoing the subsidy). As an illustration, consider households in the 1990 SIPP. In the absence of child support, 54 percent of all couple-headed families not already on welfare would gain financially if the lower-earning parent were to stop working and start receiving welfare and the higher-earning parent were replaced by an unrelated adult with identical earnings.[2] Among the households that would gain, the median gain is $892 (in 2009 dollars), a 59 percent increase in monthly household income. Among families in which neither parent had finished high school, 76 percent would gain, and among families in the bottom quarter of the income distribution, 90 percent would gain.

The enforcement of child support obligations would reduce the proportion of families standing to gain from such behavior, but only slightly. Suppose a child's father were replaced by a man who is himself a father. Part of the new man's earnings would go toward child support. That loss of household income would not be offset by support received from the child's own father because, in most states, a family's welfare benefit is reduced by the amount of support payments collected on its behalf. Even accounting for child support, however, the proportion of households that would gain by swapping the father for another man remains large. Extending the example in the preceding paragraph, if one assumes child support to be set at 15 percent of the absent parent's income (and if one makes the conservative assumption of perfect compliance), the fraction of parents without high school diplomas that would gain by separating falls from 76 percent to 65 percent.[3]

While one could refine these calculations by taking more factors into account (such as work expenses, the upward-sloping time profile of many workers' earnings, the imputed value of Medicaid benefits, or the incomplete enforcement of child support obligations), the qualitative conclusion would remain: during the years examined here, welfare's implicit subsidy for replacing fathers with other men was large enough to be relevant for many couples, especially low-income couples.

So far, we have made the case only that welfare contains an indirect but large incentive for multiple-father fertility in that it subsidizes family breakup, and thus may pave the way for the entry of an unrelated man. If that man then has a child with the children's mother, a multiple-father family is created.

A second and different incentive for multiple-father family formation is created if the welfare system treats a multiple-father family more generously than a one-father family. Because it is possible to receive benefits for some members of the household even though other members are ineligible, an unmarried mother is entitled to higher benefits if her cohabitor is the father of only some of her children than if he is the father of all of them. The children whose father is absent remain categorically eligible for benefits. Thus, among unmarried cohabitors, welfare favors the multiple-father family over the one-father family. Other things equal, however, there is no advantage to increasing the number of fathers beyond two.

III. Methodology and Data

Like many researchers who study couples' behaviors – cohabitation, marriage, divorce, fertility, and child support – we eschew a structural model in favor of a reduced-form approach. There is a large theoretical literature on the behavior of couples, but the models are not generally very amenable to testing because so few of the theorized constructs are observable. A mother's decision to have children with more than one man depends, for example, on the men's qualities as companions or (step)fathers and on the mother's own qualities as a companion, but such qualities can be hard to measure. To take another example, in bargaining models of couple behavior, each partner has an implicit threat point. Identifying that threat point is difficult for the other partner, let alone the researcher. The fact that each partner's behavior is strategic, that is, determined partly by the other's actions, poses additional difficulty. Moreover, people are forward-looking, making it hard to correlate conditions in one period with relationship status observed in a later period.

As for other, more observable factors, the numerous theoretically plausible interactions among them make it hard to predict their net influence on the likelihood of multiple-father fertility. Factors such as a mother's net gain from working, the expected value of child support, the expected value of welfare benefits, the probability of being able to find a male companion in the future, the expected income of that future partner, and the cost of housing may have different implications for women's behavior than for men's.

Some of those factors – e.g., housing costs, benefit levels, or conditions in the local labor market – can be measured fairly directly. A mother's net gain from working cannot be observed (because of the endogeneity of her labor supply decision), but her *potential* net gain from working depends on her personal characteristics, conditions in the local labor and housing markets, and the level of welfare benefits. Similarly, her probability of finding another male partner cannot be measured directly, but depends on her personal characteristics and on local demographics.

At present, the existing research on multi-partnered fertility is almost entirely descriptive rather than theoretical. Researchers have documented some broad associations. Men who have been incarcerated, for example, are more likely to have children by multiple women (Guzzo and Furstenberg 2007a, Logan, Manlove, Ikramullah, and Cottingham 2006, Mincy 2002, and Carlson and Furstenberg 2006). Welfare recipients have higher rates of multi-partnered fertility (Meyer, Cancian, and Cook 2005), as do men and women with less education, African-Americans, women who have their first births as teens, and women with nonmarital first births

(Carlson and Furstenberg 2006).

The only theoretical model of multiple-father fertility of which we are aware is a model of marriage and fertility proposed by Willis (1999, S33). In that model, if "females are in excess supply and have sufficiently high incomes, a marriage market equilibrium may exist in which children are born within marriage to high-income parents, whereas in low-income groups men father children by multiple partners outside of marriage." This study can be seen as a test of the implications of the Willis model. We focus on the correlation between multiple-father fertility and a specific form of income, namely, welfare benefits, and control for the local sex ratio to allow for the possibility that "females are in excess supply."

Given the many theoretically plausible interactions among observable factors, we estimate two variants of the following reduced-form model of multiple-father fertility (MFF):

(1) $$MFF_{ismt} = \beta_0 + \beta_1 b_{st} + \beta_2 Z_i + \beta_3 S_{st} + \beta_4 MSA_{mt} + u_{ist}$$

The subscripts *i*, *s, m* and *t* denote woman *i* living in state *s* in MSA *m* in year *t*, and *MFF* is an indicator of multiple-father fertility. The first variant is a binary logit model of whether the mother has children by more than one man. The second is a multinomial logit model with six possible outcomes, each corresponding to a different combination of fertility status and marital history. The parameter of particular interest is β_1, the coefficient on the state- and year-specific AFDC benefit, b_{st}. Our measure of benefits is the maximum benefit for a four-person household (in 2009 dollars).[4] Z_i, MSA_{mt}, and S_{st} are vectors of maternal, MSA (Metropolitan Statistical Area), and state characteristics, respectively.

Our main data source is the U.S. Census Bureau's Survey of Income and Program Participation (SIPP), a series of large-scale longitudinal surveys designed to be nationally representative. The unit of analysis is a mother with resident children. Taking one observation per family per panel and pooling 9 panels (1985-1988, 1990-1993, and 1996) yields observations on 47,653 mothers. This sample, weighted so as to be nationally representative, is the basis for our descriptive statistics.[5]

Although SIPP includes respondents from every state, state of residence is identified only for respondents in the 42 most populous states.[6] Thus, our regression results are for a slightly smaller sample that excludes the roughly three percent of respondents from the masked states and two percent of observations with missing data values. One model, one that includes MSA-level explanators, is based on a sample that is nearly fifty percent smaller, because SIPP masks the MSA for such a large fraction of respondents.[7]

The advantage of SIPP for this research is that, in each survey except that of 1989, the precise relationship of each person in a household to every other person in that household is recorded. This matrix of household relationships is the basis of our measure of multiple-father fertility.[8]

To determine the number of fathers represented among a mother's children, we use a two-step procedure. First, we identify, for each woman, every household member who is listed as her biological child. Then, restricting our attention to the interrelationships among her children, we count the number of occurrences of "Full sibling" and of "Half-sibling," which enables one to infer the number of fathers represented among those children.

This method for measuring multiple-father fertility has several limitations. One is that, if a mother has more than 5 children present, we can infer only whether she has children by more than one man; we cannot, for instance, tell a case of two fathers from a case of three. For this reason, we exclude from our sample the 1.1 percent of mothers with six or more children present (a small group of mothers with a higher-than-average likelihood of having children by more than one man). A second limitation is that some of a mother's children may not be living with her; they may live with their fathers or other relatives, or may have grown up and moved out. A third is that some mothers, especially younger ones, have not yet finished having children. Each of these measurement issues causes us to underestimate the incidence of MFF.

There is also the possibility of a small degree of attrition bias in our sample, as the detailed household relationships are collected not in the first of each panel but the second.[9] We do not anticipate this to be a serious bias, however, as only four months elapse between the first and second waves.[10]

Maternal characteristics controlled for in the analysis are age, age at the time her oldest resident child was born, ethnicity, education, and marital history. The gap between the mother's current age and her age at the birth of the oldest child in her household tells us how long she has been exposed to the possibility of MFF. Her current age (together with a set of year dummies) controls for country-wide trends in norms or expectations that could affect women's fertility decisions. For instance, the stigma of welfare receipt or of multi-partnered fertility may have been different in 1985 than in 1996. Similarly, the rise between 1985 and 1996 in political hostility toward welfare spending may have altered expectations about the future value of welfare benefits. The education variables indicate whether the parent has less or more than a high school education. Our two controls for marital history are dummy variables for "Never married" and "Ever divorced."

Because welfare benefits are set at the state level, it is important to control for other state-level factors, socioeconomic as well as political, that could be correlated with both family structure and benefit levels. We include three state-level controls besides the welfare benefit: annual measures of the state divorce rate, per capita income, and the strictness of child support enforcement.

A state's divorce rate could affect the incidence of multiple-father fertility, directly as well as indirectly. For the population as a whole, the most common route to multiple-partner fertility is divorce; in our sample, 65 percent of mothers who have children by more than one man have been divorced. A state's divorce rate also reflects the social norms and demographic composition of its population, themselves determinants of welfare policies and fertility choices.

We control for a state's per capita income for similar reasons. A richer state may choose higher benefit levels, other things equal, and higher incomes may also be associated – positively or negatively – with higher levels of multiple-partner fertility. In the Willis model, for instance, it is the conjunction of high enough incomes for women with a low sex ratio that results in multiple-partner fertility.

The significance of child support enforcement is that, where enforcement is stricter, welfare's incentive for mothers and fathers to separate is smaller, other things equal. Between 1985 and 1996, nearly all states stepped up their enforcement efforts, due in large part to provisions in the 1988 Family Support Act (FSA) that enhanced states' ability and motivation to collect child support, particularly on behalf of children on AFDC. The level of enforcement efforts varied considerably across states and over time (see, for example, Nixon 1997; Bitler 2001; Plotnick, Ku, Garfinkel, and McLanahan 2004; Huang *et al.* 2005). Like many other researchers, we proxy the strictness of enforcement with the annual ratio of the number of paternities established by a state's child support enforcement agency to the number of non-marital births in that state.[11]

The 1988 FSA mandated other changes in states' welfare programs as well – such as stricter work requirements and more extensive job training programs – but the impact of those provisions was greatly muted by the grossly inadequate funding provided for implementing them (Rushefsky 2002). The FSA also required that every state have an AFDC-UP program but, as noted earlier, the expansion of AFDC-UP does not alter the incentive under scrutiny here.

MSA-level characteristics included in the analysis are the cost of housing, the sex ratio, the median wage for men and for women, and the male unemployment rate. Our measure of housing costs is the 40th percentile of local rents, known as the Fair Market Rent (FMR) and published annually for each MSA by the U.S. Department of Housing and Urban Development. For sex ratio, we assume a racially segmented partner market and compute, from 1990 Census data, the race-specific ratio of employed men per woman in each MSA.[12] Median male and female wages and the male unemployment rate are based on the earnings and employment status of all SIPP respondents aged 22 to 50 in that MSA. Following Moffitt (2000), median wages enter the analysis as the sum of the male and female wages and also as the male-to-female ratio of wages.

IV. Empirical results

Descriptive statistics

In our SIPP sample, 8.4 percent of mothers in our SIPP sample have children by more than one man (see Table 1). Breaking down that 8.4 percent, we see that 7.59 percent have children by two men, 0.73 percent have children by three men, and 0.08 percent have children by 4 or more men. In other words, mothers with children by one man outnumber those with children by two men more than tenfold, who in turn outnumber mothers with children by three men more than tenfold, the latter themselves outnumbering mothers with children by four men or more nearly tenfold.

[TABLE 1 TO GO HERE]

Another contrast is between families that report receiving public assistance and families that do not. As Table 2 shows, the frequency of multiple-father families is markedly higher among mothers receiving aid (17.1 percent) than among other mothers (7.4 percent), in keeping with the findings of Meyer *et al.* (2005) and Guzzo and Furstenberg (2007).

[TABLE 2 TO GO HERE]

Ethnicity is strongly correlated with the likelihood of multiple-father fertility (see Table 3). The incidence of MFF is highest among black mothers (14.0 percent). Hispanic mothers have a slightly higher rate (10.0 percent) than white mothers (7.3 percent), and Asian mothers have the lowest rate (3.6 percent).

[TABLE 3 TO GO HERE]

The likelihood of multiple-father fertility also depends a great deal on the age at which a woman has her first child. Mothers who had their first child when they were fifteen or younger have an overall rate of MFF of 25.7 percent, five times the rate for mothers who were 25 or older (see Table 4).

[TABLE 4 TO GO HERE]

Looking at the data year by year, there appears to be a slight increase in the rate of multiple-father fertility between 1985 and 1996. Whether mothers are grouped by ethnicity, by education, by metropolitan status, or by marital history, the trend is upward, albeit slightly, in nearly every subgroup of mothers.[13] Consequently, inter-group differences appear relatively constant over the period. The rate of MFF among African-American mothers is consistently double or triple that of other mothers, for example (see Figure 1). The rate among mothers with more than a high school education is generally half the rate of mothers with a high school education and a quarter of that of high school dropouts (see Figure 2). Urban mothers have consistently lower rates of MFF than other mothers (see Figure 3), although the urban-rural difference is likely to be understated in SIPP data.[14] Mothers who have been divorced have an MFF rate that is generally 5 percentage points higher than mothers who have never married, who in turn have a rate that is generally 11 percentage points higher than that of mothers who are still in their first marriage (see Figure 4).

[FIGURES 1, 2, 3 AND 4 TO GO HERE]

Regression analysis

To explore the possible correlation between welfare benefits and multiple-father fertility while controlling for the interrelatedness of factors such as education, ethnicity, welfare receipt, age at first birth, and marital history, we turn to regression analysis. We consider two classes of model: a binary logit model of whether a mother has children by more than one man (Table 5), and a multinomial logit model in which the different outcomes correspond to specific combinations of multiple-father fertility and marital history (Tables 6 and 7). For ease of interpretation, we report not logit coefficients or odds ratios, but rather a variable's marginal effect on the probability that an observation is in a particular category, computed at the mean of that variable (see Appendix Table 1 for the variables' means and standard deviations).

The binary model's results suggest that welfare benefits are significantly and positively correlated with the probability of multiple-father fertility, but that the effect is very small. Other things equal, a $100 increase in the welfare benefit is associated with an increase in the probability of MFF of less than a fifth of a percentage point. The multinomial logit results offer a more nuanced picture, suggesting that the observed correlation between welfare and MFF is largely the result of a correlation between welfare benefits and the probability that a mother has a non-marital first birth.

[TABLE 5 TO GO HERE]

Table 5 reports estimates for three binary logit models. The first includes only the welfare benefit, some of the mother's personal characteristics, and a set of year dummies. The second model adds some state characteristics as well, and the third adds some MSA-level explanators. The reader will note that this third model is based on a much smaller sample, that its goodness of fit is slightly lower despite the additional explanators, and that it yields a nearly identical estimate of the correlation with welfare. For these reasons, we limit our comments here to the second model and, in all subsequent regressions, exclude the MSA-level explanators.[15]

The correlation between benefit levels and multiple-father fertility in the first model in Table 5 just misses the five-percent threshold of statistical significance. In the second model, however, which controls for three other state characteristics, the correlation is highly significant, positive, and small.[16] A $100 increase in the welfare benefit (a 13.4 percent increase relative to the mean benefit) is associated with a 0.16 percentage point increase in the probability that a mother has children by more than one man. Even the maximum effect size is a mere quarter of a percentage point. This is a small effect, in absolute terms and relative to the baseline rate of 8.6 percent.

The marginal effects reported in Table 5 for a mother's personal characteristics are consistent in sign with the unadjusted differences shown in Tables 1-4 and Figures 1-4 but are considerably smaller. For instance, more education still reduces the likelihood of multiple-father fertility, but the roughly 6 percentage point difference we saw in Figure 2 between mothers with a high school education and those with post-secondary education falls to a 2 percentage point difference after one controls for other factors. Similarly, the difference in MFF rates between African-American mothers and white mothers falls from 6.7 percentage points to 2.5 percentage points. The difference between white and Hispanic mothers loses significance, as does the difference between white and Asian mothers.

The regression results in Table 5 point to the two principal pathways to multiple-father fertility: early childbearing, and childbearing by divorced mothers. The three factors with the largest effects on a mother's probability of multiple-father fertility are (in descending order) having a child before the age of 16, having been divorced, and having a child at age 16 or 17.

Having a first birth at 15 or younger raises the probability by 23 percentage points, being divorced raises it by 14 percentage points, and having a child at 16 or 17 raises it by 11 points.

Early childbearing and having been divorced are not mutually exclusive behaviors, to be sure, but they overlap little enough that they can be viewed as distinct paths to multi-partnered fertility.[17] This suggests that we might sharpen our analysis by differentiating among three varieties of multiple-father fertility: that occurring in the context of divorce and remarriage, that occurring in the context of nonmarital childbearing and later marriage, and that occurring entirely outside marriage. In this study, we emphasize the possibility that welfare's eligibility rules contribute to multiple-father fertility by creating an incentive for mothers to live with men other than the father of their children. Those rules, however, strongly favor unmarried cohabitation over marriage. The implication is that welfare may have much more connection with the multiple-father fertility that occurs outside marriage than with that which arises in the context of divorce and remarriage. The results of the multinomial logit analysis reported in Table 6 (for the whole sample) and Table 7 (for the African-American subsample) are consistent with this hypothesis.

The six outcomes in the multinomial logit model correspond to six combinations of fertility and marital history. We categorize mothers as (a) having children by either one man or by more, and (b) having been divorced (once or more), being in their first marriage, or never having married. (For estimation purposes, the base category is mothers who are still in their first marriage and have children by only one man; three-fifths of mothers are in this category). Overall, 8.4 percent of mothers have children by multiple men. Roughly two-thirds of those mothers have been divorced (5.5 percent of the sample) and the other third have either never married (1.2 percent) or are in their first marriage (1.7 percent).

The multinomial logit estimates in Table 6 suggest that any contribution by the U.S. welfare system to multiple-father fertility is via an increased likelihood of nonmarital births. For mothers who have been divorced, the correlation between welfare benefits and MFF turns out to be insignificant (column 5). The three outcomes for which the correlation with benefits is statistically significant all involve nonmarital births. Benefits are positively correlated with being a never-married mother with children by one man (column 1), with being a never-married mother with children by multiple men (column 3), and with being a mother still in her first marriage but with children by multiple men (column 4).[18]

[TABLE 6 TO GO HERE]

These estimated effects, while statistically significant, are very small. A $100 increase in benefits corresponds to a 0.04 percentage point increase in the likelihood that a woman has never married but has children by multiple men, and a 0.11 percentage point increase in the likelihood that she is in her first marriage and has children by more than one man. Even the upper bounds of these effects are only 0.06 and 0.15 percentage points, respectively. To put these effects into perspective, consider that the marginal effect of having a baby before age sixteen is over 50 times larger.

Sensitivity analyses

As a check on the plausibility of the estimates in Tables 5 and 6, we repeat the analysis using a number of different subsamples. The results are qualitatively similar across all of them, suggesting that the welfare benefit effect is not a mere statistical artifact. When we limit our sample, for example, to non-African American mothers (or, more narrowly, to white mothers), the regression results (not shown) closely resemble the results for the entire sample. When we restrict our attention to African-American mothers, the results (shown in Table 7) are qualitatively similar, but the point estimates are slightly larger. As in the overall sample, benefits are uncorrelated with MFF among mothers who have been divorced, but are correlated with the three outcomes that entail nonmarital births. For instance, a $100 increase in benefits corresponds to a 0.23 percentage point increase in the likelihood that a never-married African-American mother has children by multiple men. While that effect is five times larger than the effect found for other mothers, five times a tiny effect is still a small effect.

[TABLE 7 TO GO HERE]

A potentially informative sensitivity test is to repeat the analysis on a sample from which the more educated mothers are excluded. The rationale for excluding them is that they are the mothers for whom one would expect the level of welfare benefits to be the least relevant to fertility decisions. If the correlations with welfare benefits are not spurious, one would expect them to become more pronounced when the most educated mothers are excluded, and to get larger still when the sample is restricted to the least educated. This is, in fact, what we find, in every instance (results not shown). The estimated effects for the least educated mothers (those who did not finish high school) are roughly twice the size of those reported in Tables 5 and 6 for the entire sample. If we restrict the sample even further to African-American mothers who did not finish high school, the largest estimated effect is an increase of only seven-tenths of a percentage point in the likelihood of MFF.

Another sensitivity check is to limit the sample to mothers whose oldest child is six years old or younger. The purpose of this restriction is to limit the amount of time that may have elapsed between the year when a mother had a child by a second man and the year the family is surveyed, so as to increase the pertinence of the level of welfare benefits in the survey year. Again, if the estimated welfare effects are not spurious, one would expect them to be larger for this subsample, and again, they are.

As a test specifically of the Willis (1999) model, in which multiple-father fertility occurs when mothers' incomes are high enough and the male/female sex ratio is low enough, we restrict the sample to mothers for whom MSA is known (because our sex ratio variable is at the MSA-level) and add an indicator for cases in which both conditions are met. We experiment with a variety of thresholds for defining "high enough" income and a "low enough" sex ratio, but in none of these regressions is the coefficient on that indicator significant.

Concerned that Alaska and Hawaii, because of their unusually high costs of living, could unduly influence the results, we exclude observations from those two states and repeat the analysis. The impact on the estimates is virtually imperceptible.

In a final test, we add state fixed effects, to control for any omitted state characteristics correlated with both benefits and multiple-father fertility. In both models, this raises explanatory power only slightly (the pseudo-R^2 rises by about .005), and the marginal effect of the welfare benefit (and of each of the other state-level variables) becomes insignificant (results not shown). A possible interpretation is that the estimates in Tables 5 and 6 are indeed spurious, generated by some correlation at the state level between norms conducive to multiple-father fertility and norms favoring higher benefits. Before leaping to that conclusion, however, we make two points. First, recall that the estimated effects are larger when the sample is limited to less educated mothers, or African-American mothers, or mothers with younger children, and that benefits were correlated with MFF for mothers who had had nonmarital births but not for mothers who had been divorced; it is hard to reconcile such patterns with the hypothesis that benefit effects are spurious and the result of omitted state characteristics. Second, in these particular data, the fixed effects estimates constitute a weak test of that hypothesis, because within-state variation in the real AFDC benefit is so much smaller than between-state variation. As Figure 5 shows, in the 15 largest states (which account for about 85 percent of welfare recipients), real benefits were not highly variable between 1985 and 1996. In most states, the real benefit declined slowly over time as a result of inflation, and only in a few instances did benefits change enough to affect a state's ranking. In short, because AFDC benefits are so highly collinear with the state fixed effects, any real AFDC effects, particularly very small ones, are likely to be masked by the inclusion of state effects. In the end, however, it makes little difference whether one inclines to the state fixed effects models or the other models; the biggest estimated effects that we find are still strikingly small.

[FIGURE 5 TO GO HERE]

V. Conclusion

This study explores the possible connection between the U.S. welfare system and multi-partnered fertility during the decade preceding TANF. Welfare's eligibility rules, intended to steer public aid to children in the poorest households, contain financial inducements for a mother to live with somebody other than her child's father. This incentive is large enough to be relevant for a broad swath of the population. In our sample of mothers with resident children – a pooled cross-section of the 1985-1996 Surveys of Income and Program Participation – roughly half of all couples not already on welfare would have gained financially if the children (or at least some of them) had been fathered by someone other than the man of the house, with the median gain exceeding 50 percent of current household income.

Despite the size of the incentive and its relevance for so many families, our data suggest that the behavioral response it generates is minimal. In the model and sub-population yielding our largest estimated effect, a $100 increase in the welfare benefit corresponds to an increase in the probability of multiple-father fertility of only seven-tenths of a percentage point. When state fixed effects are included, the correlation between benefits and the likelihood that a mother has children by more than one man goes away altogether. In short, the size of the estimated effect ranges from zero to seven-tenths of a percentage point, depending on one's choice of model and subsample. We cannot reject the hypothesis that the level of welfare benefits is a determinant of the prevalence of multiple-father fertility, but our results suggest that, if welfare's influence is real, it is so slight as to be negligible.

This analysis focuses on the pre-TANF era, but the findings remain relevant for the present. If the large incentives built into the pre-TANF welfare system did not contribute significantly to the phenomenon of multi-partnered fertility, then it is unlikely that our current welfare system, with its time-limited benefits and stricter work requirements, is an important contributor to the rise in multi-partnered fertility. Policymakers would do well to keep this in mind when they consider sacrificing other policy objectives -- such as income adequacy for the poorest children -- in the hopes of improving incentives in the welfare system.

References

Bitler, M. (2001). The effects of child support enforcement on non-custodial parents' labor supply. Unpublished paper.

Carlson, M. J. and Furstenberg, F.F. (2006). The prevalence and correlates of multi-partnered fertility among urban U.S. parents. *Journal of Marriage and Family*, 3, 718-732.

Center for Economic and Policy Research. (2009). *SIPP Talking Points*. Retrieved October 21, 2010, from http://www.ceprdata.org/savesipp/talkingpoints.pdf

Cooksey, E. C. and Craig, P. H. (1998). Parenting from a distance: The effects of paternal characteristics on contact between nonresidential fathers and their children. *Demography*, 2, 187-200.

Evenhouse, E. and Reilly, S. (2005). Father taxes and father absence. Unpublished paper.

Guzzo, K. B. and Furstenberg, F. F. (2007a). Multi-partnered fertility among American men. *Demography*, 3, 583-601.

Guzzo, K. B. and Furstenberg, F. F. (2007b). Multi-partnered fertility among young women with a nonmarital first birth: Prevalence and risk factors. *Perspectives on Sexual and Reproductive Health*, 1, 29-38.

Harknett, K. and Knab, J. (2007). More kin, less support: Multi-partnered fertility and kin support among new mothers. *Journal of Marriage and the Family*, 1, 237-253.

Huang, C. C., Mincy, R. B., and Garfinkel, I. (2005). Child support obligations and low-income fathers. *Journal of Marriage and Family*, 5, 1213-1225.

Klerman, L. V. (2007). Multi-partnered fertility: Can it be reduced? *Perspectives on Sexual and Reproductive Health*, 1, 56–59.

Logan, C., Manlove, J., Ikramullah, E. and Cottingham, S. (2006). Men who father children with more than one woman: A contemporary portrait of multiple-partner fertility. *Child Trends Research Brief 2006-10*. Washington, DC: Child Trends.

London, R. A. (1998). Trends in single mothers' living arrangements from 1970 to 1995: Correcting the Current Population Survey. *Demography*, 1, 125-131.

Manning, W. and Smock, P. (1999). New families and nonresident father-child visitation. *Social Forces*, 1, 87-116.

Manning, W., Stewart, S. D. and Smock, P. (2003). The complexity of fathers' parenting responsibilities and involvement with nonresident children. *Journal of Family Issues*, 5, 645-667.

Margolis, R. and Mykyta, L. (2008). Multi-partnered fertility and relationship stability. Unpublished paper.

Meyer, D. R., Cancian, M. and Cook, S. T. (2005). Multiple-partner fertility: Incidence and implications for child support policy. *Social Service Review*, 4, 577-601.

Mincy, R. (2002). Who Should Marry Whom? Multipartnered Fertility Among New Parents. Working Paper 02-03-FF. Center for Research on Child Wellbeing, Princeton University, Princeton, NJ.

Mincy, R. and Huang, C. C. (2001). 'Just get me to the church': Assessing policies to promote marriage among fragile families. Fragile Families Working Paper 2002-02-FF. Center for Research on Child Wellbeing, Princeton University, Princeton, NJ.

Moffitt, R. A. (2003). The Temporary Assistance for Needy Families program. In R. Moffitt (Ed.), *Means-tested transfer programs in the U.S.* (pp. 291-365). Chicago: University of Chicago Press/National Bureau of Economic Research.

Moffitt, R. A. (2000). Welfare benefits and female headship in U.S. time series. Department of Economics Working Paper 434, Johns Hopkins University.

Moffitt, R. A., Reville, R. T. and Winkler, A. E. (1994). State AFDC rules regarding the treatment of cohabitors. *Social Security Bulletin*, 4, 26-33.

Moffitt, R. A., Reville, R. T. and Winkler, A. E. (1998). Beyond single mothers: Cohabitation, marriage, and the U.S. welfare system. *Demography*, 3, 259-278.

Nixon, L. (1997). The effect of child support enforcement on marital dissolution. *Journal of Human Resources*, 1, 159-181.

Plotnick, R. D., Ku, I., Garfinkel, I. and McLanahan, S. S. (2004). Better child support enforcement: Can it reduce teenage nonmarital childbearing? *Journal of Family Issues*, 5, 634-657.

Rendall, M. S., Clarke, L., Peters, H. E., Ranjit, N. and Verropoulou, G. (1999). Incomplete reporting of men's fertility in the United States and Britain: A research note. *Demography*, 1, 135-144.

Roberts, P. (2008). The implications of multiple-partner fertility for efforts to promote marriage in programs serving low-income mothers and fathers. Brief No. 11 (March) in *Couples and Marriage Series*. Washington, DC: Center for Law and Social Policy.

Rolston, H. (2000). Office of the U.S. Assistant Secretary for Planning and Evaluation. Telephone conversation with author.

Rushefsky, M. E. (2002). *Public Policy in the United States: At the Dawn of the Twenty-First Century*, 3rd edition. Armonk, NY: M.E. Sharpe.

Turney, K. and Carlson, M. J. (2010). Multi-partnered fertility and mental health among fragile families. Fragile Families Working Paper WP10-03-FF, Center for Research on Child

Wellbeing, Princeton University, Princeton, NJ.

Willis, R. J. (1999). A Theory of Out-of-Wedlock Childbearing. *Journal of Political Economy*, 6 (Part 2), S33-64.

Table 1

Number of fathers represented among a mother's children, by number of children

Number of children living with mother	**One father**	**Two fathers**	**Three fathers**	**Four or more fathers**	**Total**
1	15,429				*15,429*
2	17,352	1,729			*19,081*
3	7,360	1,278	208		*8,845*
4	2,218	460	98	23	*2,799*
5	613	101	35	15	*765*
Percentage of all mothers	***91.59***	***7.60***	***0.73***	***0.08***	
Totals	*42,972*	*3,567*	*341*	*38*	*46,919*

Notes: Unit of analysis is a mother with resident children. Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels, and are weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded.

Table 2

Number of fathers represented among a mother's children, by receipt of public aid

Number of fathers	**Family not receiving aid**	**Family receiving aid**	**Totals**
1	38,909	4,065	*42,794*
2	2,893	672	*3,566*
3	198	143	*342*
4+	17	21	*38*
MFF rate (%)	***7.4***	***17.1***	
Totals	*42,018*	*4,901*	*46,919*

Notes: Unit of analysis is a mother with resident children. Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels, and are weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded.

Table 3

Number of fathers represented among a mother's children, by race/ethnicity

Number of fathers	**White**	**Black**	**Hispanic**	**Asian**	**Totals**
1	31,977	5,766	3,863	1,369	*43,853*
2	2,346	785	387	47	*3,668*
3	178	122	36	4	*346*
4	6	23	4	0	*34*
5	0	5	0	0	*5*
MFF rate (%)	***7.3***	***14.0***	***10.0***	***3.6***	
Totals	*34,507*	*6,702*	*4,290*	*1,420*	*46,919*

Notes: Unit of analysis is a mother with resident children. Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels, and are weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded.

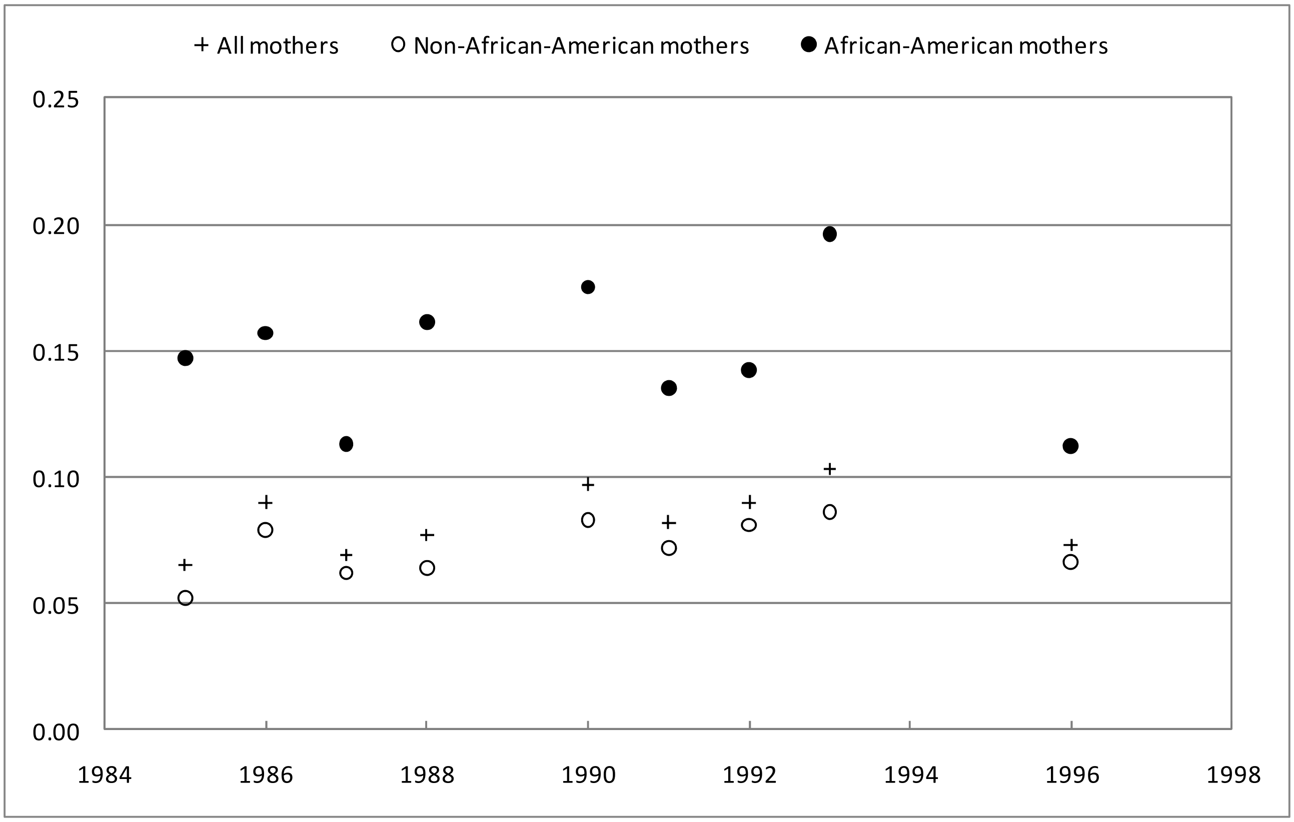
Table 4

Number of fathers represented among a mother's children, by age at first birth

Number of fathers	**Aged 15 or younger**	**Aged 16-17**	**Aged 18-19**	**Aged 20-24**	**Aged 25 or older**	**Totals**
1	279	1,595	3,698	13,909	23,494	*43,853*
2	75	308	641	1,373	1,167	*3,668*
3	18	42	65	121	95	*346*
4	3	6	11	7	7	*34*
5	0	0	0	4	1	*5*
MFF rate (%)	***25.6***	***18.3***	***16.2***	***9.8***	***5.1***	
Totals	*375*	*1,950*	*4,416*	*15,413*	*24,764*	*46,919*

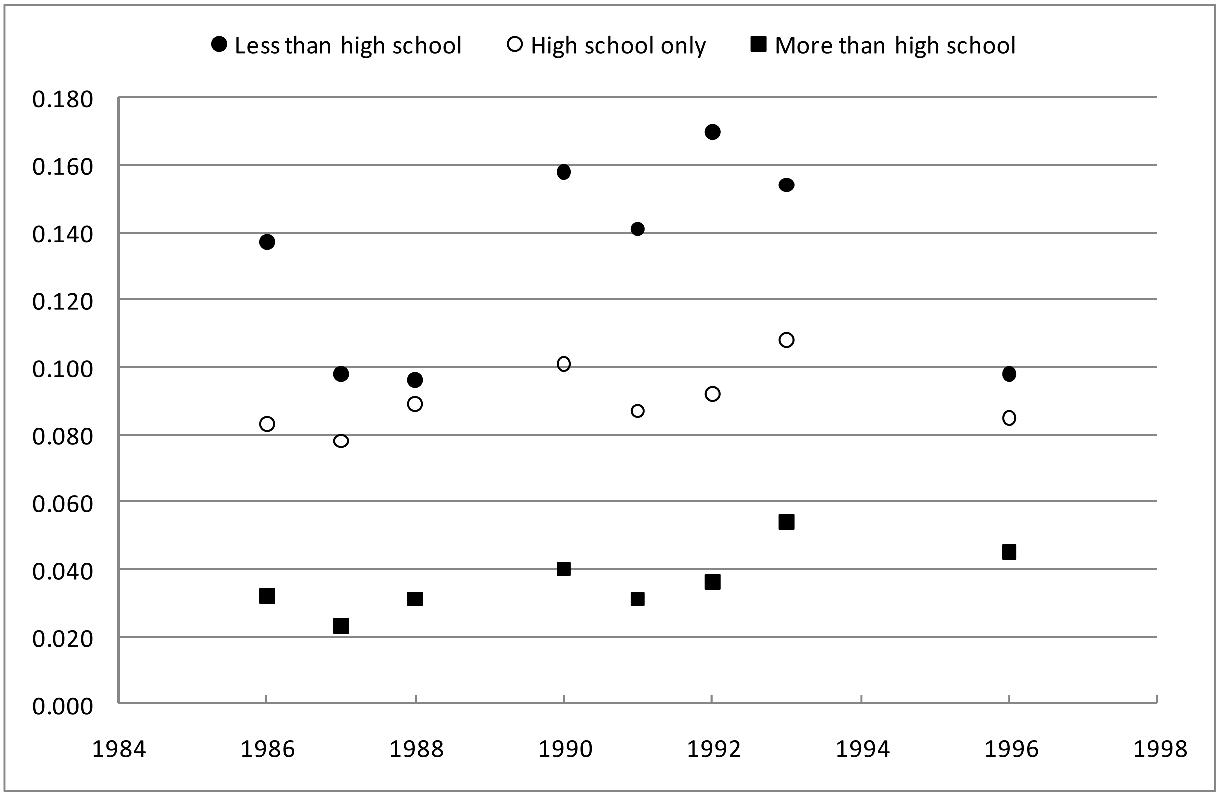
Notes: Unit of analysis is a mother with resident children. Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels, and are weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded

Figure 1
Fraction of mothers with children by more than one man, by ethnicity



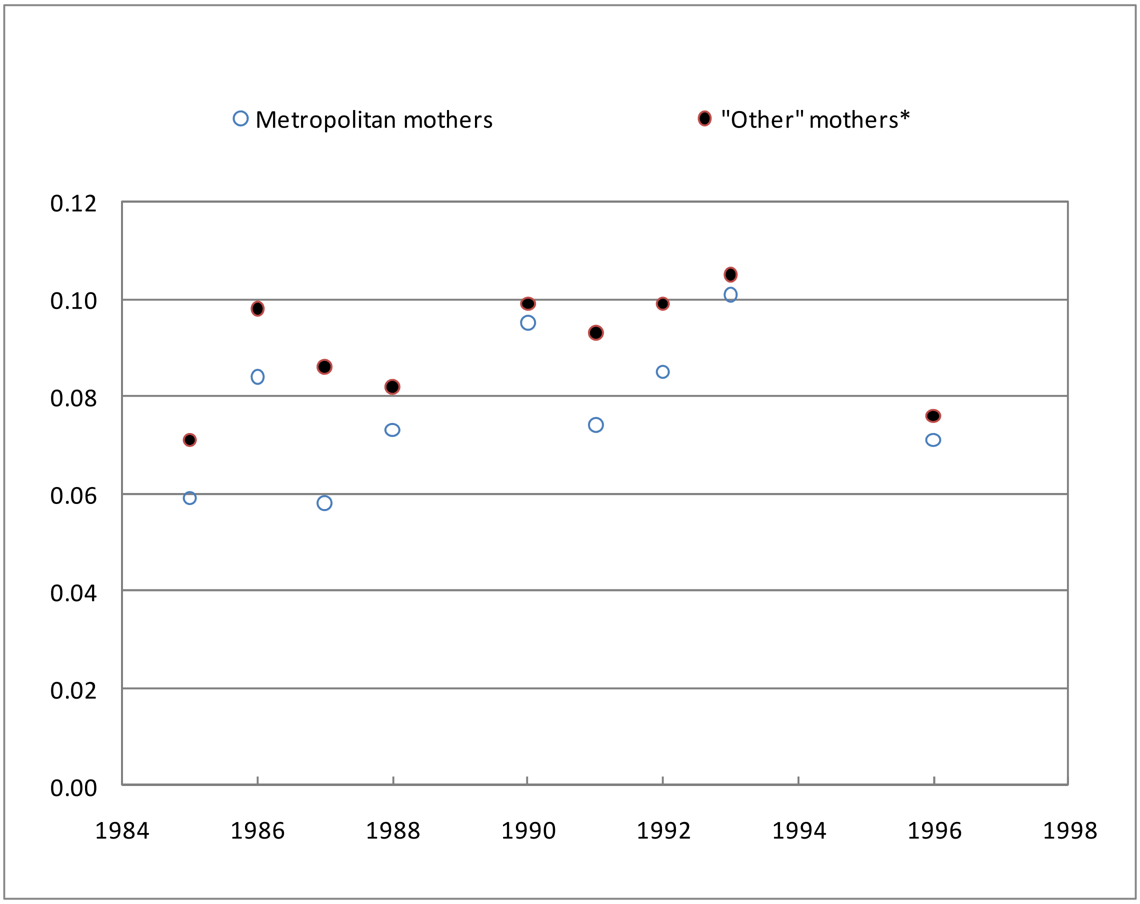
Notes: Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels. Data weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded.

Figure 2
Fraction of mothers with children by more than one man, by education



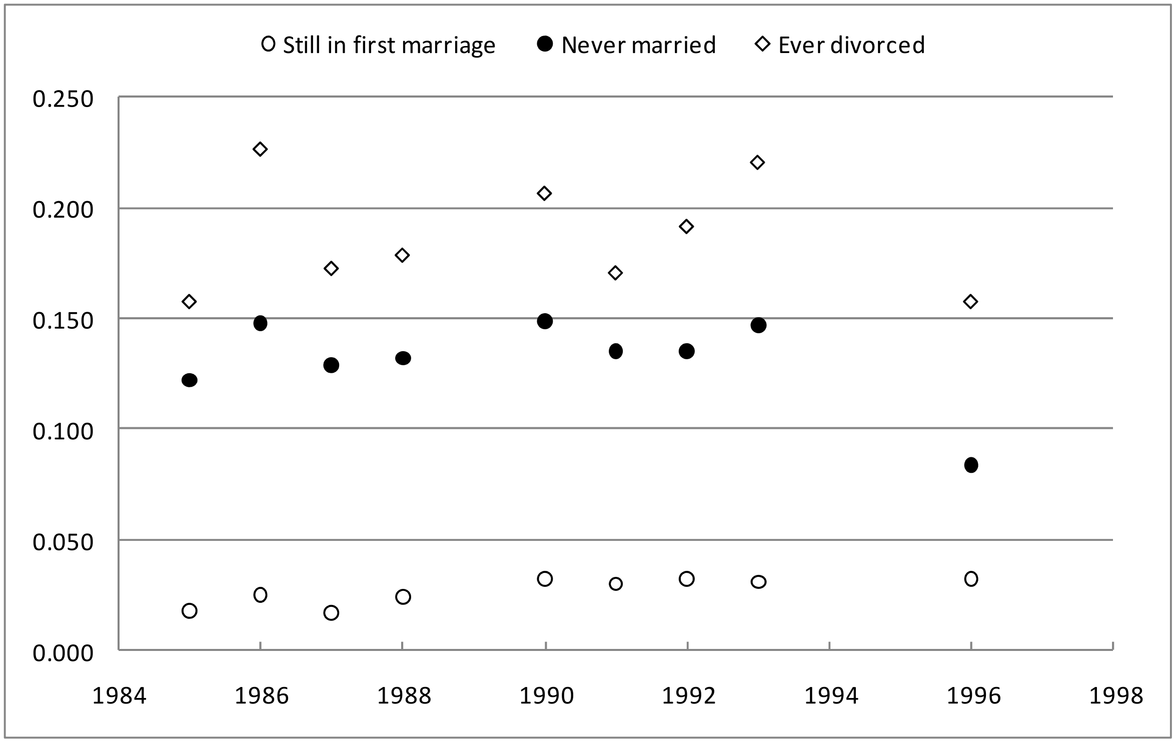
Notes: Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels. Data weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded.

Figure 3
Fraction of mothers with children by more than one man, by metropolitan status*



*Notes: Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels. Data weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded. *"Other" is not synonymous with non-metropolitan. SIPP randomly masks the MSA for a large share of respondents in smaller metropolitan areas, so "other" combines those respondents with genuine non-metropolitan respondents.*

Figure 4
Fraction of mothers with children by more than one man, by marital history



Notes: Data are from the 1985-1988, 1990-1993, and 1996 SIPP panels. Data weighted to be nationally representative. Mothers with 6 or more children present—1.06 percent of SIPP mothers—are excluded.

Table 5. Binary logit models: "One father" versus "Two or more fathers"

	Model 1 **Mother's personal characteristics**	*Model 2* **State characteristics added**	*Model 3* **MSA characteristics added**
AFDC benefit *($100s of 2009 $)*	.0006 *(.051)*	**.0016** *(.000)*	**.0015** *(.020)*
------- *Personal characteristics* -------			
Black	**.0236** *(.000)*	**.0250** *(.000)*	**.0213** *(.005)*
Hispanic	- .0044 *(.160)*	.0059 *(.071)*	**.0131** *(.001)*
Asian	- .0093 *(.156)*	-.0081 *(.243)*	-.0070 *(.393)*
Never married	**.0772** *(.000)*	**.0786** *(.000)*	**.0704** *(.000)*
Has been divorced	**.1461** *(.000)*	**.1448** *(.000)*	**.1329** *(.000)*
Aged 15 or less at birth of oldest child	**.2444** *(.000)*	**.2328** *(.000)*	**.2243** *(.000)*
Aged 16-17 at birth of oldest child	**.1139** *(.000)*	**.1100** *(.000)*	**.1146** *(.000)*
Aged 18-19 at birth of oldest child	**.0858** *(.000)*	**.0830** *(.000)*	**.0917** *(.000)*
Aged 20-24 at birth of oldest child	**.0271** *(.000)*	**.0265** *(.000)*	**.0317** *(.000)*
Less than high school education	**.0130** *(.000)*	**.0131** *(.000)*	**.0095** *(.006)*
More than high school education	**- .0196** *(.000)*	**-.0194** *(.000)*	**-.0205** *(.000)*
Age *(years)*	**.0155** *(.000)*	**.0153** *(.000)*	**.0143** *(.000)*
Age squared	**-.0002** *(.000)*	**-.0002** *(.000)*	**-.0002** *(.000)*
-------- *State characteristics* ----------			
Intensity of child support enforcement *(0 - 1.2)*		**.0100** *(.042)*	.0056 *(.473)*
Per capita income *($1,000s of 2009 $)*		**-.0006** *(.017)*	.0001 *(.814)*
Divorce rate		**.0031** *(.004)*	.0032 *(.054)*
-------- *MSA characteristics* -----------			
Rent level *(40th percentile, in $100s of 2009 $)*			**-.0032** *(.008)*
Race-specific sex ratio *(employed men/woman)*			-.0094 *(.622)*
Sum of median male and median female wage			-.0002 *(.381)*
Ratio of median male to median female wage			-.0025 *(.340)*
Male unemployment rate			-.0002 *(.495)*
Year dummies	Yes	Yes	Yes
Proportion of obs in the "2+ fathers" category	0.086	0.086	0.083
Sample size	45,436	43,308	22,592
Pseudo-R^2	0.161	0.162	0.158

Notes: Omitted outcome category is one father. Table reports marginal change in probability of 2 or more fathers for 1-unit change in variable (p-value of underlying logit coefficient in parentheses). Bold font denotes significance at 5-percent level or better. Omitted racial/ethnic category is non-Hispanic white. Omitted marital status is "still in first marriage." Omitted education category is high school diploma. "Oldest child" means "oldest child currently in household." Family structure data from 1985-88, 1990-93, and 1996 SIPP surveys.

Table 6. Six-outcome multinomial logit model

	One father and…		Multiple fathers and…		
	…never married	…divorced (once or more)	…never married	…still in first marriage	…divorced (once or more)
AFDC benefit *($100s, in 2009 $)*	**.0031** *(.000)*	-.0012 *(.914)*	**.0004** *(.000)*	**.0011** *(.000)*	.0002 *(.368)*
------- *Personal characteristics* -------					
Black	**.2192** *(.000)*	**.0091** *(.000)*	**.0273** *(.000)*	**.0241** *(.000)*	**.0001** *(.000)*
Hispanic	**.0202** *(.000)*	**-.0479** *(.000)*	**.0017** *(.008)*	.0012 *(.773)*	-.0034 *(.182)*
Asian	-.0007 *(.115)*	**-.1353** *(.000)*	.0027 *(.243)*	.0027 *(.811)*	**-.0349** *(.000)*
Aged 15 or less at birth of oldest child	-.0080 *(.111)*	**.0130** *(.003)*	**.0223** *(.000)*	**.0925** *(.000)*	**.1600** *(.000)*
Aged 16-17 at birth of oldest child	**-.0072** *(.000)*	**.1174** *(.000)*	**.0092** *(.000)*	**.0445** *(.000)*	**.1233** *(.000)*
Aged 18-19 at birth of oldest child	**-.0017** *(.000)*	**.0829** *(.000)*	**.0081** *(.000)*	**.0297** *(.000)*	**.0935** *(.000)*
Aged 20-24 at birth of oldest child	-.0025 *(.387)*	**.0390** *(.000)*	**.0029** *(.000)*	**.0105** *(.000)*	**.0272** *(.000)*
Less than high school education	**.0282** *(.000)*	**.0290** *(.000)*	**.0054** *(.000)*	**.0060** *(.000)*	**.0098** *(.000)*
More than high school education	**-.0275** *(.000)*	**-.0639** *(.000)*	**-.0028** *(.000)*	**-.0081** *(.017)*	**-.0239** *(.000)*
Age *(years)*	**-.0149** *(.000)*	**.0325***(.000)*	-.0000 *(.283)*	**.0014** *(.000)*	**.0231** *(.000)*
Age squared	**.0001***(.000)*	**-.0003** *(.000)*	**-.0000** *(.035)*	**-.0000** *(.001)*	**-.0003** *(.000)*
-------- *State characteristics* ----------					
Intensity of child support enforcement	-.0019 *(.598)*	-.0173 *(.215)*	.0007 *(.459)*	.0056 *(.064)*	.0054 *(.430)*
Per capita income *($1,000s of 2009 $)*	-.0001 *(.585)*	.0008 *(.264)*	-.0000 *(.530)*	-.0002 *(.085)*	-.0003 *(.372)*
Divorce rate	**-.0052** *(.000)*	**.0230** *(.000)*	-.0001 *(.975)*	**.0008** *(.039)*	**.0066** *(.000)*
Number of observations in category	*3,728*	*10,694*	*540*	*787*	*2530*
Proportion of observations in category	*0.083*	*0.235*	*0.012*	*0.017*	*0.055*
Sample size		44,791			
Pseudo-R^2		0.130			

Notes: Model also includes year dummies. Omitted outcome is "One father, still in first marriage" (59.8 percent of observations). Table reports change in outcome's probability for a 1-unit change in each variable (p-value of the underlying logit coefficient in parentheses). Bold font indicates significance at the 5-percent level or better. Omitted education category is "High school education." Family structure data are from the 1985-88, 1990-93, and 1996 SIPP surveys.

Table 7. Six-outcome multinomial logit model, African-American subsample

	One father and …		Multiple fathers and …		
	…never married	…divorced (once or more)	…never married	…still in first marriage	…divorced (once or more)
AFDC benefit *($100s, in 2009 $)*	**.0155** *(.000)*	-.0030 *(.104)*	**.0023** *(.002)*	**.0018** *(.022)*	.0002 *(.088)*
------- *Personal characteristics* -------					
Aged 15 or less at birth of oldest child	-.1139 *(.310)*	-.0632 *(.160)*	**.1744** *(.000)*	**.1748** *(.000)*	**.0377** *(.004)*
Aged 16-17 at birth of oldest child	**-.0654** *(.001)*	**.0683***(.000)*	**.1059** *(.000)*	**.0567** *(.000)*	**.0465** *(.000)*
Aged 18-19 at birth of oldest child	**-.0459** *(.002)*	**.0615** *(.000)*	**.0777** *(.000)*	**.0490** *(.000)*	**.0280** *(.000)*
Aged 20-24 at birth of oldest child	-.0411 *(.379)*	**.0523** *(.000)*	**.0323** *(.000)*	**.0175** *(.002)*	**.0212** *(.000)*
Less than high school education	**.1235** *(.000)*	**-.0276** *(.009)*	**.0275** *(.000)*	-.0036 *(.118)*	**.0008** *(.013)*
More than high school education	**-.0938** *(.000)*	**.0154** *(.011)*	**-.0273** *(.000)*	**-.0166** *(.002)*	**-.0022** *(.045)*
Age *(years)*	**-.0673** *(.000)*	**.0459***(.000)*	.0020 *(.156)*	**.0086** *(.001)*	**.0259** *(.000)*
Age squared	**.0006** *(.001)*	**-.0004** *(.000)*	-.0001 *(.058)*	**-.0001** *(.003)*	**-.0003** *(.000)*
-------- *State characteristics* ----------					
Intensity of child support enforcement	**.0438** *(.043)*	-.0010 *(.201)*	**.0252** *(.006)*	.0111 *(.173)*	.0054 *(.275)*
Per capita income *($1,000s of 2009 $)*	-.0008 *(.245)*	**-.0027** *(.050)*	.0009 *(.612)*	-.0005 *(.224)*	**-.0003** *(.013)*
Divorce rate	**-.0278** *(.008)*	.0057 *(.994)*	.0048 *(.164)*	.0045 *(.293)*	.0066 *(.402)*
Number of observations in category	*1,943*	*1,548*	*331*	*243*	*370*
Proportion of observations in category	*0.295*	*0.239*	*0.049*	*0.038*	*0.058*
Sample size		6,425			
Pseudo-R^2		0.113			

Notes: Model also includes year dummies. Omitted outcome is "One father, still in first marriage" (32.2 percent of observations). Table reports change in outcome's probability for a 1-unit change in each variable (p-value of the underlying logit coefficient in parentheses). Bold font indicates significance at the 5-percent level or better. Omitted education category is "High school education." Family structure data are from the 1985-88, 1990-93, and 1996 SIPP surveys.

Appendix Table 1. Sample characteristics

Variable	**Mean**	**Std dev**	**N**
AFDC benefit *(in $100s)*	7.46	3.25	45,436
African-American	0.149		46,919
Hispanic	0.091		46,919
Asian	0.030		46,919
Never married	0.097		46,919
Has been divorced	0.289		46,919
Aged 15 or less at birth of first child	0.008		46,919
Aged 16-17 at birth of first child	0.041		46,919
Aged 18-19 at birth of first child	0.092		46,919
Aged 20-24 at birth of first child	0.329		46,919
Less than a high school education	0.167		46,919
More than a high school education	0.232		46,919
Age *(years)*	34.6	8.11	46,919
Year=1985	0.018		46,919
Year=1986	0.169		46,919
Year=1987	0.073		46,919
Year=1988	0.066		46,919
Year=1990	0.127		46,919
Year=1991	0.072		46,919
Year=1992	0.112		46,919
Year=1993	0.101		46,919
Year=1996	0.262		46,919
Intensity of child support enforcement *(Paternities established per nonmarital birth)*	0.413	0.255	46,919
State per capita income *(in $1,000s)*	30.8	5.37	46,919
State divorce rate *(Divorces per 1,000 residents)*	3.97	2.02	46,919
Rent level *(40th percentile, in $100s of 2009 $)*	8.77	2.53	27,409
Sex ratio *(employed men per woman)*	0.805	0.235	27,409
Sum of median male and median female wage *($ per hour)*	47.3	14.9	27,409
Ratio of medial male to median female wage	1.90	0.78	27,409
Male unemployment rate *(percent)*	8.23	4.94	27,409

Notes: All dollar figures in 2009 dollars. Data weighted to be nationally representative.

Figure 5. 1985-1996 trends in AFDC benefits (15 largest states)

[1] Welfare's treatment of stepparent income may have factored into fertility behavior, but during the years studied here, that treatment was the same across the states. Prior to 1981, states had discretion in the matter, and generally chose to ignore much or all of a stepfather's income; after 1981, all states were required to take that income into account. The impact was immediate: between May 1981 and May 1982, stepfather households fell from 6.6 percent of the AFDC caseload to 3.4 percent (Rolston 2000). The federal waivers received by several states in the 1990s that let them disregard a larger share of stepparent income—to reduce the marriage disincentive—were not tantamount to a restoration of the pre-1981 regime; the disregards were typically modest, temporary, or both (Brown 1995).

[2]In other words, the potential gain is measured as the difference between the state-specific AFDC/ Food Stamp benefits package for a family of four and the lower of a couple's two incomes.

[3]In addition, during the 1985-1996 period, most states allowed AFDC families to keep the first $50 of child support and reduced their monthly benefit, dollar for dollar, for amounts collected above $50. For low-income parents, this worked to reduce the importance of child support obligations as a deterrent to separation.

[4]The AFDC benefit data are courtesy of Robert Moffitt.

[5] SIPP includes numerous weights. The weights used in this analysis were the monthly person weights, which permit the analyst to use all available data for a given month. The person weight corresponds to the inverse probability of selection, with adjustments for subsampling within clusters, for non-response, and for movers. For more detail, see http://www.census.gov/sipp/weights.html (last accessed January 18, 2010).

[6] The masking of states varies slightly by survey year (with the exception of Maine and Vermont, which are combined in every panel). In the 1996 SIPP, for example, North Dakota, South Dakota, and Wyoming were combined, whereas in the 1985 survey, Iowa was combined with North and South Dakota, and Wyoming was combined with Alaska, Idaho, and Montana.

[7] In states with small nonmetropolitan populations, the Census Bureau randomly recodes a substantial share of SIPP's metropolitan respondents as nonmetropolitan; see any SIPP User Guide for more information.

[8]These household relationship codes appear to be highly reliable. While a small fraction of them – between 0.6 and 1.2 percent, depending on the survey year – are imputed, they should not be considered missing data and hence a potential source of bias. Analysis of the 1996 survey (the first year in which SIPP recorded the type of imputation) reveals that 100 percent of the instances of imputed relationship code were logical imputations rather than statistical imputations, that is, they were based on good data for the household contained elsewhere in the survey instrument.

[9]The exception is the 1985 panel, in which household relationship detail was not collected until the fourth wave.

[10] In the 1996 SIPP, for instance, attrition after three years was 26 percent (CEPR 2009).

[11] This choice reflects the priority placed by the 1988 Family Support Act on paternity establishment, particularly for children on welfare. States were penalized if they failed to establish paternity in a given proportion of the children born to mothers on welfare, and the federal government bore 90 percent of the states' associated laboratory costs.

[12]SIPP indicates residence in 60 to 80 MSAs and roughly 20 CMSAs (grouped MSAs). Fair Market Rent (FMR) and sex ratio values for a CMSA were obtained by averaging the values for its constituent MSAs.

[13] More striking than subgroup time trends are the large changes in the MMF rate from one year to the next in almost every subgroup's time series. This volatility raises the question of whether all SIPP panels are equally representative of the U.S. population, a question well outside the scope of this study.

[14] For a substantial portion of respondents in smaller MSAs, SIPP masks the MSA information by recoding the respondent as non-metropolitan, thereby contaminating the non-metropolitan subsample, and biasing downward estimates of urban-rural differences.

[15] The sample becomes nearly 50 percent smaller with the addition of MSA-level explanators because so many respondents either live outside metropolitan areas or, as mentioned earlier, have their MSA masked by SIPP.

[16] The welfare coefficient becomes significant when any one of the state-level variables is added individually, as well as when all three are added at once.

[17] Among mothers in our sample who have been divorced and have children by more than one man, fewer than one in twelve had a child before age 18.

[18] We cannot offer an exact figure, but are making what we consider a reasonable assumption, namely, that the large majority of mothers in the latter category – mothers who are in a first marriage but have children by multiple men – had a nonmarital birth from a relationship that preceded their marriage.