The Effect of Marital Breakup on the Income Distribution of Women with Children

Abstract

Having a female firstborn child significantly increases the probability that a woman's first marriage breaks up. Recent work has exploited this exogenous variation to measure the effect of divorce on economic outcomes, and has concluded that divorce has little effect on women's mean household income. However, using a Quantile Treatment Effect methodology (Abadie et al. 2002) we find that divorce widens the income distribution: it increases the probability that a woman has very low or very high household income. It appears that some women successfully generate income through child support, welfare, combining households, and increased labor supply after divorce, while others are markedly unsuccessful. Thus, although divorce has little effect on mean income, it nonetheless increases poverty and inequality. These findings imply that divorce has important welfare consequences.

1. Introduction and Motivation

The poverty rate for single mothers fell substantially from 1974 to 2005. Over the same period, the poverty rate for married mothers remained virtually unchanged. Given these facts, one might assume that women with children are less likely to be poor than they were 30 years ago. But in fact the overall poverty rate for women with children rose over this period, from .120 to .143 (Figure 1).

A clue to resolving this puzzle may be found in the fact that both divorce and single parenthood have greatly increased over the past several decades throughout the developed world. In the United States, the proportion of mothers who are single rose from about 16 percent in 1974 to over 26 percent in 2005. It appears that in the absence of this trend, overall poverty rates would have decreased, rather than increased.

Current political discussions commonly assume that marriage has causal economic effects on women and children. In particular, recent welfare legislation encourages marriage as a method of increasing income and reducing the need or eligibility for welfare. The federal government spends $150 million per year on the Healthy Marriage Promotion Program (DHHS 2007), part of its overall poverty-reduction plan; in addition, over 40 state governments now make their own official efforts to support marriage (Dion 2005).

Despite the assumptions made in popular debate, a causal relationship between marriage preservation and poverty reduction has not been demonstrated. Previous research has established the need for instrumental variable (IV) analysis in order to identify any effect of family status on women's outcomes (Becker, Landes, and Michael 1977; Becker 1985; Angrist and Evans 1998; Gruber 2000). Recently, first-born child sex has emerged as an instrument for marital status, facilitating estimates of causality (Morgan and Pollard 2002, Lundberg and Rose 2003, Dahl and Moretti 2004). Using that instrument, Bedard and Deschenes (2005) conclude that "IV results cast doubt on the widely held view that divorce causes large declines in economic status for women" (p. 411). While it is true that the negative correlation between mean income and divorce appears to be driven by selection, we demonstrate in this paper that a conclusion based on the mean effect of divorce is misleading. In fact, an examination of the impact of divorce on the entire income distribution supports the view that divorce greatly increases the odds of poverty.

This paper also uses the sex of the first-born child as an instrument for marital breakup and then conducts IV analysis to separate the causal effects of divorce from its well-known correlations. With data from the 1980 U.S. Census, we document that having a female first-born child slightly, but robustly, increases the probability that a woman's first marriage breaks up. In our sample, the likelihood that a woman's first marriage is broken is 0.63 percentage points higher if her first child is a girl, representing a 3.7 percent increase from a base likelihood of 17.2 percentage points. A discussion of the mechanisms that drive this instrument, and detailed investigations into its validity and robustness, can be found in the working paper version of this paper (CITE), in Bedard and Deschenes (2005) and in Dahl and Moretti (2004).

Unlike these previous papers, however, we use a Quantile Treatment Effect (QTE) estimation strategy (Abadie et al. 2002) that allows us to look at the impact of marital breakup throughout the income distribution. Using this technique, we find that marital breakup significantly affects the household income distribution. In particular, marital breakup dramatically increases the probability that a mother will end up with a very low income or a very high income.

It follows logically that divorce increases the probability of living in a household without other earners. In fact, we estimate that breakup of the first marriage significantly increases the likelihood that a woman lives in a household with less than $5000 of annual income from others—the likelihood rises from just over 5 percent for those whose first marriage is intact to nearly 50 percent for those whose first marriage breaks up. However, women can and do respond to income loss from divorce by combining with other households, through paths including remarriage or moving in with a roommate, sibling, or parents. Moreover, women further compensate through private (e.g. alimony and child support) and public (e.g. welfare) transfers, and by increasing their own labor supply. Since, further, divorce reduces family size as well as income, the net effect of the husband's departure on the household's income-to-needs ratio is ex ante ambiguous. In other words, it may be possible for a woman to entirely offset the loss of her husband's income so that her material well-being is undiminished.

When examining the entire income distribution, however, we find that these responses, although substantial, are frequently insufficient to prevent poverty. While virtually none of the women influenced by our instrument who remain in their first marriage are in poverty, nearly a quarter of those who divorce are in poverty. In fact, IV results suggest that divorce causes an increase in poverty that is about as large as that suggested by the observed correlation between divorce and poverty.

We show that the lack of effect on income at the mean (Bedard and Deschenes 2005) comes from the fact that the cumulative distribution functions of income of women who remain married and those who divorce cross each other—we find evidence that some women who divorce, rather than moving lower in the income distribution, move towards the top of the income distribution, possibly due to re-marriage outcomes or to moving in with their parents. For example, at the 95th percentile, women who divorce have household incomes over twice as high as those who remain married. We conclude that breakup of the first marriage increases the variance of income, even though there is no significant effect on the mean. Because breakup of the first marriage leads some women to have higher incomes as well as leading more women to be poor, estimates that focus on the mean do not detect the dramatic effect that divorce has on the income distribution. That is, while divorce does not affect average income, it does exacerbate inequality and poverty. We view this result as part of a growing body of evidence on the importance of analyzing the effects of policy and social changes over the entire distribution of outcomes (Bitler, Gelbach, and Hoynes 2006; Neumark, Schweitzer, and Wascher, 2004; Blank and Schoeni 2003), rather then merely at the mean.

After discussing our findings, we consider the aggregate implications of the relationship we identify between divorce, poverty and inequality. In recent decades, mothers' poverty rates have failed to decline as much as the overall poverty rate. At the same time, income inequality has increased. We argue that the increase in the proportion of broken first marriages in the U.S. could help account for both the stagnation in mothers' poverty and the increase in income inequality

The paper proceeds as follows. In section 2, we describe the data and the sample we use. In section 3, we describe the estimation strategy. In section 4, we discuss the results. In section 5, we conclude.

2. Data

We use data on women living with minor children from the 5 percent 1980 Census file (Ruggles, Sobek et al. 2003), which allow us sufficient power to identify the effect of sex of the first-born child on marital breakup.[1] We limit our sample to white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. These limitations are necessary in order to create a sample for which measurement error in the sex of the observed first-born child has a classical structure. Bedard and Deschenes (2005) take a similar although not identical approach to sample limitation, and get estimates very close to ours; Dahl and Moretti (2004) show a variety of samples, the most similar of which gives first-stage estimates that are very close to ours. Further discussion of the sample construction is contained in the Appendix.

3. Estimation Framework

In many applications of the OLS framework, researchers are interested in estimating the mean effect of a regressor (e.g. the breakup of a woman's first marriage) on various outcomes (such as measures of her income). The quantile regression (QR) model of Koenker and Bassett (1978) allows one to estimate the effect of a regressor not only on mean outcomes, but on the entire distribution of outcomes. Since we are interested in the distributional effects of divorce on income, this framework is useful for our purposes.

[1] While earlier censuses have these measures, in previous decades the divorce rate was very low. Subsequent censuses, on the other hand, do not have all of the measures necessary to conduct the analysis.

The endogeneity of marital breakup poses a challenge for interpreting the results of QR as the causal impact of divorce on the income distribution. But just as 2SLS can be used to overcome problems of omitted variables bias measurement error with OLS, so can the Quantile Treatment Effect (QTE) framework of Abadie et al. (2002) be used to overcome these problems in QR.

To apply the QTE model to our research question, we define our regressor of interest, D, as a dummy for the breakup of the first marriage. We consider a woman as having her first marriage intact if she reported both that she was "currently married with spouse present" and that she had been married exactly once. We consider a woman as having her first marriage broken if she: has been married multiple times, is married but currently not living with her husband, is currently separated from her husband, is currently divorced, or is currently widowed.[2]

Our main outcomes, Y, are total others' income, defined as total household income less total own income; household income; and two measures of household income adjusted for household size. The change in others' income measures the direct effect on a woman of losing her husband's income (to the extent that the husband is not replaced by other wage earners). The change in total household income captures this direct effect but also includes the indirect effects of divorce on income: transfers from the ex-husband in the form of alimony and child support;[3] transfers from the state in the form

[2] We have run the analysis with ever-divorced, rather than first marriage broken, as the explanatory variable: in this case, widows and those separated from or not living with their first husband are coded as 0 rather than 1. Our results are not sensitive to this difference in categorization.
[3] The 1980 Census question reads: "Unemployment compensation, veterans' payments, pensions, alimony or child support, or any other sources of income received regularly… Exclude lump-sum payments such as money from an inheritance or the sale of a home."

of cash assistance; and income generated by changes in the woman's own labor supply.[4] We use two methods to adjust household income for family composition to capture the change in need that accompanies a change in household size. The first is household income as a percent of the federal poverty line; the second is household income divided by (number of adults + number of children$^{0.7}$)$^{0.7}$, a measure used in Bedard and Deschenes (2005) and referred to as "normalized household income."

Our controls, denoted by X, are a vector of pre-determined demographic variables which include age, age squared, age at first birth and an indicator for high-school dropouts.[5] While our Ordinary Least Squares (OLS) and Quantile Regression (QR) estimates of the relationship between marital breakup and income may be sensitive to their inclusion, our IV specifications (both 2SLS and QTE) are robust to controls. We think of U as representing unobserved factors such as human capital, views on gender roles, and taste for non-market work relative to market work and leisure. Finally, our instrumental variable, Z, is an indicator for having a girl as one's firstborn child.

We now consider two potential outcomes indexed against D: Y_1 is the value a woman's outcome if her marriage breaks up, and Y_0 is the outcome if the first marriage remains intact. Similarly, D_1 tells us whether a woman's first marriage is broken if her eldest child is a girl, and D_0 tells us whether her first marriage is broken if her eldest child is a boy. We are interested in the distribution of outcomes for women whose first

[4] We have also added each of these sources of income into our outcome measures sequentially. Private and public transfers account for very little income, and so adding them to income from others does not materially change the distribution of income (results available from the authors).

[5] Since we look only at women who gave birth to their first child at age 19 or later, we can reasonably assume that the decision on whether to graduate from high school is made prior to the realization of the sex of the first-born child.

marriage is broken and for women whose first marriage remains intact, and in the causal effect of divorce on the income distribution, which is the difference: $Y_1 - Y_0$.

Abadie et al. (2002) provide assumptions under which we can estimate a Quantile Treatment Effect (QTE) model:

1. Independence: (Y_1, Y_0, D_1, D_0) is jointly independent of Z given X.
2. Non-trivial assignment: $0 < P(Z = 1 \mid X) < 1$.
3. First stage: $E[D_1|X] \neq E[D_0|X]$.
4. Monotonicity: $P(D_1 \geq D_0|X)=1$.
5. For each θ ($0<\theta<1$) there exist α_θ and β_θ such that $Q_\theta(Y|X, D, D_1>D_0) = \alpha_\theta D + X'\beta_\theta$, where $Q_\theta(Y|X, D, D_1>D_0)$ denote the θ-quantile of Y given X and D for compliers.

Under these assumptions, and additional regularity requirements (see Abadie et al. 2002) we can compute a consistent estimator of α_θ and β_θ,. The first four assumptions outlined above are very similar to the assumptions Angrist and Imbens (1994) use to derive the Local Average Treatment Effect (LATE) interpretation of 2SLS,[6] and have been previously tested for the case of child sex as a predictor of divorce in Bedard and Deschenes (2005), Dahl and Moretti (2004), and the working paper version of this paper (CITE). Finally, the fifth assumption requires that we can write a linear model of quantile regression, where our primary parameter of interest is α_θ, which gives the

[6] In this setting, the first assumption implies that child sex is randomly assigned. Although we cannot formally test this assumption, we find that mothers of girls and boys aged 0 or 1 are not significantly different from each other, consistent with the hypothesis that child sex is randomly assigned in our sample. The second assumption simply states that both girls and boys account for a non-trivial fraction of births. The third assumption is that there is a significant first stage, which we demonstrate below. The fourth assumption is that there are no women for whom having a first-born girl reduces the probability of divorce. In other words, while our model requires that there are compliers (women whose first marriage breaks up only if their first child is a girl), there are no defiers (women whose first marriage breaks up only if their first child is a boy).

difference in the conditional θ-quantiles of Y_1 and Y_0 for compliers. In other words, just as IV can be used to estimate the LATE of marital breakup on income for women whose first marriage is broken if and only if they have a girl (called "compliers" in the Angrist and Imbens literature), so can QTE estimate the effect of marital breakup on compliers' income distribution.

4. Results

The first column of Table 1 gives the estimated relationship between the sex of the first-born child and marital status for our sample. We find that having a girl increases the probability of breakup of the first marriage by about 0.63 percent; due to the large sample size, this effect is measured very precisely.[7] The effect of sex of the eldest child on the breakup of the first marriage is bigger than that on being currently divorced, which is 0.20 percent, because many of the women who divorced due to having a girl subsequently remarry (an endogenous response that is not properly part of our IV analysis). It is important to keep in mind that we are not estimating the effect of being currently divorced or of residing in a mother-only household on economic outcomes. Rather, we are estimating the effect of having *ever* been divorced on current outcomes.

Cross-sectional (OLS) regressions of the relationship between marital breakup and various measures of income (shown in the second column of Table 1), which admit no causal interpretation, show that breakup of the first marriage is correlated with lower income. These relationships confirm the conventional view that women whose first

[7] We have also replicated this result using the Current Population Survey 1980, 1985, 1990, and 1995 Fertility Supplements, which identify the sex of a woman's actual firstborn child regardless of whether that child is still in the household, and hence require no sample restrictions. The CPS estimate of 0.68 percent is highly similar to that from our Census sample. Similarly, Bedard and Deschenes (2005), using a slightly different sample, derive an estimate of 0.80, with a standard error of 0.10.

marriages end are significantly worse off than women whose first marriages remain intact.

In the right-hand column, we replicate the results of Bedard and Deschenes (2005). Two-stage least squares estimates suggest that the mean causal effect of marital breakup on material well-being is quite different from the population correlation. The two-stage estimate of the effect of divorce on others' income is negative but insignificant and of much smaller magnitude than the OLS estimate. The two-stage estimate of the effect of divorce on household income level, although not significantly different from zero, is positive and significantly different from the OLS estimate. The mean effects on adjusted household income and on income as a percent of poverty are both positive, but not precisely estimated. Bedard and Deschenes interpret these results as indicating that on average there is negative selection into divorce—women who would have had low income anyway are more likely to divorce, creating a negative cross-sectional correlation between income and divorce—and that there is no significant causal effect of divorce on mean income.

Recent literature, however, has focused on and argued for the importance of examining distributional effects rather than merely effects at the mean. Bitler, Gelbach, and Hoynes (2006), Neumark, Schweitzer, and Wascher (2004), Blank and Schoeni (2003) and others claim that analysis of the effects of policy and social changes should focus on effects throughout the distribution. Moreover, in the case of marital breakup, it seems highly plausible that there are important effects on the income distribution that aren't evident at the mean. After all, divorce—and the implied withdrawal of the husband's income—represents a discrete fall in income. To the extent that women remarry, move in with parents or other relatives, or realize a high earning potential, they may end up as well or even better off financially than those who stay in their first marriage—resulting in little effect of divorce on the mean of the distribution. And those women who cannot find other sources of income could nonetheless experience large losses.

In Tables 2 and 3, the columns of results labeled "First marriage intact" report income at each decile of the distribution for those who remain married, estimated using the methods described above. Recall that in quantile treatment effects regressions, the income distributions are estimated for the specific group of women—"compliers"—who divorce or stay married in response to the sex of the first-born child. The quantile regressions represent the income distributions for the population of women as a whole. The columns labeled "Difference in income: broken – intact" report the estimated difference in income at each decile of the distribution between those who have divorced and those who stay married, which is the coefficient on marital breakup.

Standard quantile regression (QR) estimates of the difference in distribution of income by marital breakup (left-hand panel in Table 2) tell a familiar story. Divorce is associated with lower levels of income of others in the woman's household throughout the distribution; the same is true for all deciles of total household income. In cross-section, women with broken first marriages have income distributions —both household and others'— that are first-order stochastically dominated by those of women whose first marriages remain intact. That is, divorce is correlated with a shift downward in income at every point in the income distribution.[8]

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[8] Note also that the estimated income distribution for compliers whose first marriage remains intact is also uniformly lower than the income distribution for the full sample. This result, which holds in all of our

The quantile treatment effects (QTE) estimates, in the right panel of Table 2, give a more nuanced picture. We do in fact find a large effect of marital breakup on the probability of having very low income: roughly 40 percent of women whose first marriage is broken have no household income due to any other household member. While legal transfers (which include child support and mean-tested transfers) reduce the number of compliers with no income, one in six of the divorced compliers have no unearned income even after accounting for transfers—compared to virtually none of the compliers who stay married. These effects on the bottom of the distribution are similar to the naïve QR estimates.

The analysis of household income, which adds in each woman's own earnings, also shows a large effect of marital breakup on the bottom of the income distribution. Roughly one in six of those who experience marital breakup have less than $5000 in household income, compared to virtually none of those who remain married.[9]

Interestingly, the QTE estimates also show that those who divorce due to the instrument are *more* likely to have high levels of household income. These results diverge from naïve QR estimates, suggesting that in the population at large negative selection into divorce swamps these effects and makes them undetectable. The top 20 percent of women whose first marriage breaks up because of a firstborn girl have more income from others than do those who remain married. At the 95th percentile, women who divorce have more than twice as much income from others as do women who avoid

estimates, suggests that couples whose marriages break up in response to the instrument typically have a low SES.

[9] Although they cannot earn enough to make up for the loss in others' income, divorced compliers do have a very large labor supply response to the loss. Since the distributional effect on hours does not differ markedly from the mean responsiveness of about 1100 hours reported in Bedard and Deschenes (2005), a separate analysis is not included here.

divorce; the difference in the 95th percentile is statistically significant. Previous literature (Mueller and Pope 1980) finds that when divorced women remarry, their second husband is typically more educated and has a higher occupational SES score. In addition, Bedard and Deschenes (2005) argue that many divorced mothers co-reside with their parents who, due to lifecycle effects, have higher incomes than their husbands did.[10]

When women's own earnings are added in to reflect total household income, the magnitude of the effect of divorce on the top of the income distribution is even more striking, although less precise. Roughly 40 percent of women who divorce due to the instrument have higher household income than counterparts who remain married. This percentage may be higher because top-end variation in the rewards to women's increased labor supply is added to top-end variation in the income of new household members.

As shown in Figure 2, the reversal in the sign of the income gap between those who divorce and those who do not occurs above the median of the distribution for all types of income, meaning that the typical family has less income when the first marriage breaks up. However, as is often the case with income distributions, average income among compliers is greater than median income (result available from the authors); as a result, the reversal occurs at (in the case of others' income) or below (in the case of household income) the *mean* of the distribution. This statistical artifact explains why models that estimate effects at the mean, such as those in Bedard and Deschenes (2005), find little or positive effect of marital breakup on income.

[10] Bedard and Deschenes split their sample into those whose oldest child is under 12 years of age and those whose oldest child is 12 to 16; they find that the effect on others' income is important mostly for those with older children, who they argue are less likely to combine households. We, similarly, find that the negative effect of divorce is greater for those with older children (results available from the authors).

The results discussed so far ignore one important aspect of divorce—namely, it reduces family size. Thus even if a woman loses income through divorce, she may not necessarily end up worse off if there are also fewer family members to support. On the other hand, if income gains at the top come heavily through combining households, the effect may be neutralized or worse by increased household size. To estimate the effect of divorce on the ratio of income to needs, we divide each woman's total household income by the federal poverty line (FPL) for a household of that size. Doing so also tells us about eligibility for programs such as Medicaid, SCHIP, free and reduced price lunch, and child care subsidies, since eligibility cutoffs for these programs are based on FPL values such as 100% and 180%. We also use an alternative adjustment for family size used in Bedard and Deschenes (2005).

As shown in Table 3, we find that changes in family size do not fully offset the effect of marital breakup on household income. The naïve QR estimates indicate that marital breakup decreases normalized household income at all levels. At the bottom of the income distribution, the QTE estimates are again comparable to naïve QR estimates, showing that fewer than five percent of women still in their first marriage have household income below the poverty line, while nearly a quarter of those whose first marriage ended are below poverty. The bottom 40 percent of those who divorce due to the instrument have significantly lower income-to-needs ratios than their counterparts who did not divorce.

Again, however, women whose first marriage ended are more likely to have very high income-to-needs ratios—above 400 percent of the FPL—than are compliers whose first marriage remains intact. The top ten percent of the compliers who divorce have significantly higher normalized incomes than those who remain married. This reversal at the top is not picked up by naïve QR estimates. The results are highly similar for the alternative normalization of household income.

5. Discussion

Our results confirm previous findings that negative selection into divorce accounts for the observed relationship between marital breakup and lower mean income. Yet marital breakup does have a significant causal effect on the distribution of income: it increases the percent of women at the bottom and top tails of the income distribution. On net, divorce increases poverty and inequality for women with children.

Our sample is not representative of all single mothers in the US—although their first marriages are broken, many of the divorced compliers are not single when observed. In addition, the women we analyze were all married before their first child was born, so our results cannot be simply generalized to never-married mothers. Further, we are able to look only at white women; the effects of divorce may differ for women of other races. Nonetheless, examining this group of compliers has at least one significant advantage: the distribution of income for compliers who are still in their first marriage is somewhat, but not drastically, lower than the overall distribution of income in society. This characteristic suggests that the coefficients we estimate reflect effects of divorce on women who are somewhat disadvantaged and "at risk" yet not far outside the mainstream in terms of socioeconomic status.

Our results suggest that the destabilization of first marriages may have caused some of the stagnation in poverty rates of women with children over the last several decades. A back-of-the-envelope calculation suggests that the poverty rate among women with our sample characteristics was 14 percent (1.5 percentage points) higher in 1995 than it would have been had the share of broken first marriages remained at its 1980 level.[11]

Our results also suggest a relationship between the destabilization of first marriages and the widening of the income distribution. Previous literature has not emphasized the relationship between divorce and inequality, although both have increased substantially over the past three decades. If early and sustained marriages act as a form of insurance against later shocks to either partner in earning capacity, then increased divorce in a sense weakens that insurance, and may be a factor in increasing household inequality. Much of the recent literature on the causes of income inequality has focused on wage inequality and the forces that may be affecting it, such as: technology (Acemoglu 2002); the decline of labor market institutions (DiNardo, Fortin, and Lemieux 1996); and the rise of international trade. Our findings suggest that the destabilization of first marriages may also be a contributing factor to increasing income inequality.

Data Appendix

The 5 percent 1980 Census data contain several measures that allow us to analyze a woman's fertility history. These include the number of children ever born to a woman, the number of marriages, the quarter as well as year of first birth, and the quarter and year

[11] In 1995 (the most recent year for which we have data on marital history from the CPS), the poverty rate for women with our sample characteristics who are still in their first marriage was 8.7 percent, while the rate for women who have ever divorced was 21.7 percent. Thus if divorce had stayed at its 1980 prevalence (17.2 percent ever divorced), the overall poverty rate in this sample would be: 0.172*0.217 + (1-0.172)*0.087 = 0.109, or 10.9 percent. Because the prevalence of divorce rose to 28.6 percent, the poverty rate became: 0.286*0.217 + (1-0.286)*0.087 = 0.124, or 12.4 percent. Thus the increase in divorce may potentially have caused an increase of 1.5 points, or 13.8 percent, in the poverty rate for women with our sample characteristics.

of first marriage. This information permits us to identify the sex of the first-born child for most women, although not for women whose eldest child has left the household.

A substantial drawback of using cross-sectional data is the fact that we can only observe the sex of the oldest child who resides with the mother, whereas ideally we would want to observe the sex of the firstborn child. It is important that we create a sample of women for whom measurement error in the sex of the observed first-born child has a classical structure. To that end, we attempt to restrict the sample to those women observed with all their biological children. We do so in order to limit the risk that our results will be affected by differential attrition of boys and girls. In particular, we are concerned that boys are differentially more likely to end up in the custody of their fathers in the event of marital breakup. This pattern could lead to endogeneity of our instrument if the sample were left uncorrected. If, in the event of divorce, fathers keep the sons and mothers keep the daughters, there will be a spurious positive correlation in the overall sample between marital breakup and the eldest *observed* child being a girl.

To address this issue, we exclude from the sample any woman for whom the number of children ever born does not equal the number of children living with her. If a mother lives with stepchildren or adopted children in a number that exactly offsets the number of her own children that are not living with her, this rule will fail to exclude her. We therefore further minimize the possibility of including women who have non-biological children "standing in" for biological children by including only women whose age at first birth is measured as between 19 and 44.

Limiting our sample to women who live with all their children reduces the risk that differential custody rates could bias our result, but it does not eliminate this risk altogether, because we could still be more likely to include divorced women with two girls than those with one boy and one girl or those with two boys. We have tested this hypothesis, however, and found that mothers living with all their children and mothers in the overall population are equally likely to be observed with a girl as the eldest child; this suggests that sex of the eldest child is not a major determinant of living with all of one's children (results available from the authors).

We limit our sample to women whose first children were born after their first marriage, since breakup of the first marriage is our focus. If instead we included out-of-wedlock births, we would be concerned that the sex of the first child affected selection into the first marriage (Lundberg and Rose 2003). To the extent that people could learn the sex of the child before it was born and thereby select into "shotgun marriages," we may still have selection into first marriage. However, ultrasound technology was not yet widely used in 1980 (Campbell 2000), so this threat is not of particular concern. In addition, because we can only identify the beginning of the first marriage and not the end, we may include some women whose first child was born after the breakup of the first marriage. While this would weaken the first stage of our estimation, it would not bias our results.

We look only at mothers whose eldest child is a minor, since those who still live with their adult children may be a select group. Further, since girls are differentially more likely to leave home early (at ages 17 and 18), we restrict the sample to mothers whose eldest child is under age 17. Fourth, we limit our sample to white women because black women are more likely to have girls, which necessitates analyzing them separately; unfortunately, the Census does not provide enough data for a strong first stage when restricting the sample to African-Americans only. Finally, we leave out any woman whose first child was a twin, both because different-sex twins would complicate our instrument and because twins increase the number of children a woman has.

Table A shows summary statistics for our full sample relative to the overall population of women with minor children. Our sample is quite similar to the overall population except in terms of age and marital status. The women in our sample are younger than average, consistent with the requirement that a woman's eldest child is under 17. The women in our sample are also slightly less likely to be divorced, both because they are younger and because we require that they have custody of all children. And of course, unlike the overall population, women in our sample cannot be never-married. On other characteristics, however, the two groups differ little: women in our sample have slightly more education and household income than the overall population and work and earn slightly less.

In summary, the limitations we place on the sample are designed to create a group of women for whom we can measure the sex of the firstborn child with only classical measurement error. These restrictions weaken the power of our first stage, but we believe this compromise is necessary in order to minimize concerns about endogeneity of our instrument.

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Table 1. Mean Regressions

	First stage	Second stage	
Dependent variable	Instrument: first child is a girl	OLS	2SLS
First marriage is broken	0.0063		
	(0.0010)		
Currently divorced	0.0020		
	(0.0007)		
Others' income		-9241	-1041
		(42.57)	(5178)
Household income		-5577	6548
		(43.41)	(5570)
Income/Federal Poverty Line		-59	27
		(0.44)	(53)
Income/Normalized Household Size		-1635	3104
		(18.34)	(2354)

NOTE: N=619,499. The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and whose first birth was a single birth. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Poverty is calculated using 1980 Census codes that range in value from 1 to 501. Normalized income is windsorized at zero. Income is in 1980 dollars. 2SLS regressions use broken first marriage as the first-stage dependent variable. Robust standard errors are in parentheses.

Table 2a. Cumulative Distribution of Others' Income by Marital Status

	Quantile Regressions			Quantile Treatment Effects		
Percentile of distribution	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact
5th	5005	-5005	-5715	4205	-4205	-4205
	(25)	(25)	(33)	(4525)	(4525)	(4554)
10th	8005	-8005	-9418	6310	-6310	-6310
	(18)	(18)	(27)	(2947)	(2945)	(3035)
20th	11430	-11430	-12464	9005	-9005	-9005
	(17)	(17)	(22)	(2471)	(2488)	(2469)
30th	14005	-14005	-13616	10580	-10580	-10580
	(17)	(17)	(28)	(2203)	(2217)	(2269)
40th	16190	-14755	-12970	12010	-11685	-10878
	(16)	(103)	(61)	(1954)	(1972)	(3948)
50th	18505	-10465	-9806	14005	-6500	-6249
	(16)	(71)	(67)	(2628)	(5569)	(5256)
60th	20640	-8110	-7680	15210	-3245	-3643
	(17)	(58)	(57)	(2431)	(3731)	(4290)
70th	23500	-6605	-6254	17010	-1900	-1851
	(21)	(62)	(52)	(3360)	(4261)	(4276)
80th	27015	-6010	-5440	19030	-25	-103
	(29)	(68)	(58)	(4437)	(5650)	(4848)
90th	34410	-6665	-4885	22005	3000	3102
	(56)	(105)	(93)	(6582)	(8535)	(8679)
95th	44305	-9130	-4866	24210	33025	31324
	(128)	(180)	(163)	(9969)	(16270)	(24731)

NOTE: N=619,499. The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and whose first birth was a single birth. All the regressions (except where indicated) include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Income is in 1980 dollars. QTE regressions use broken first marriage as the first-stage dependent variable. Robust standard errors in parentheses.

Table 2b. Cumulative Distribution of Household Income by Marital Status

	Quantile Regressions			Quantile Treatment Effects		
Percentile of distribution	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact
5th	7185	-5100	-5628	6510	-5045	-5215
	(27)	(43)	(41)	(3431)	(3849)	(3985)
10th	10160	-6150	-6736	8605	-5100	-5346
	(19)	-(33)	(38)	(2779)	(3198)	(3035)
20th	14285	-7075	-7441	11185	-4675	-5003
	(19)	(41)	(36)	(2683)	(3463)	(3239)
30th	17205	-7200	-7512	13340	-3835	-4484
	(18)	(36)	(37)	(2593)	(3390)	(3201)
40th	19910	-7100	-7154	15165	-3045	-3635
	(15)	(43)	(42)	(2345)	(3382)	(3616)
50th	22010	-6095	-6363	17015	-1710	-2243
	(17)	(49)	(48)	(2768)	(4006)	(4027)
60th	24770	-5525	-5338	19005	-85	-435
	(18)	(54)	(53)	(3107)	(4704)	(4393)
70th	27810	-4800	-4258	20795	2080	1691
	(23)	(61)	(57)	(3412)	(5179)	(5104)
80th	31950	-3945	-3163	23020	5195	4769
	(29)	(72)	(69)	(4567)	(6776)	(6367)
90th	40005	-3995	-1954	26435	13690	12793
	(44)	(112)	(107)	(7091)	(12237)	(12998)
95th	50010	-5000	-1747	29515	32635	31185
	(70)	(182)	(177)	(10942)	(19271)	(26205)

NOTE: N=619,499. The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and whose first birth was a single birth. All the regressions (except where indicated) include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Income is in 1980 dollars. QTE regressions use broken first marriage as the first-stage dependent variable. Robust standard errors in parentheses.

Table 3a. Cumulative Distribution of (Income/Federal Poverty Line) by Marital Status

	Quantile Regressions			Quantile Treatment Effects		
Percentile of distribution	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact
5th	97.00	-67.00	-66.98	101.00	-67.00	-68.43
	(0.37)	(0.64)	(0.62)	(39.51)	(45.59)	(46.73)
10th	138.00	-78.00	-78.74	128.00	-74.00	-75.26
	(0.26)	(0.46)	(0.54)	(41.32)	(44.35)	(40.77)
20th	194.00	-90.00	-81.18	171.00	-90.00	-85.76
	(0.28)	(0.58)	(0.54)	(37.21)	(40.53)	(42.56)
30th	234.00	-85.00	-79.01	207.00	-95.00	-86.39
	(0.24)	(0.59)	(0.54)	(35.10)	(43.96)	(44.16)
40th	270.00	-82.00	-75.23	238.00	-89.00	-78.13
	(0.21)	(0.60)	(0.56)	(34.97)	(48.63)	(47.30)
50th	304.00	-76.00	-71.20	262.00	-67.00	-57.04
	(0.26)	(0.69)	(0.61)	(33.64)	(60.38)	(56.59)
60th	341.00	-70.00	-64.88	284.00	-25.00	-20.80
	(0.23)	(0.66)	(0.66)	(31.09)	(68.12)	(65.92)
70th	386.00	-66.00	-56.31	307.00	27.00	27.41
	(0.37)	(0.82)	(0.75)	(30.75)	(75.60)	(70.02)
80th	446.00	-58.00	-41.12	329.00	97.00	93.94
	(0.42)	(1.08)	(0.85)	(35.54)	(84.45)	(81.63)
90th	501.00	0.00	-5.01	356.00	145.00	136.86
	(0.03)	(0.09)	(0.28)	(44.43)	(44.64)	(57.67)
95th	501.00	0.00	0.00	375.00	126.00	125.00
	(0.04)	(0.09)	(0.03)	(61.03)	(61.01)	(65.19)

NOTE: N=619,499. The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and whose first birth was a single birth. All the regressions (except where indicated) include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Poverty is calculated using 1980 Census codes that range in value from 1 to 501. QTE regressions use broken first marriage as the first-stage dependent variable. Robust standard errors in parentheses.

Table 3b. Cumulative Distribution of (Income/Normalized HH Size) by Marital Status

	Quantile Regressions			Quantile Treatment Effects		
Percentile of distribution	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact	Income when first marriage is intact	Difference: broken-intact	Difference w/controls: broken-intact
5th	3008	-1901	-1948	3579	-2886	-2865
	(11)	(21)	(20)	(1355)	(1572)	(1610)
10th	4348	-2223	-2247	4266	-2607	-2608
	(9)	(16)	(18)	(1059)	(1281)	(1297)
20th	6012	-2381	-2246	5250	-2401	-2289
	(8)	(19)	(17)	(925)	(1206)	(1239)
30th	7307	-2272	-2109	6015	-1873	-1826
	(7)	(19)	(17)	(880)	(1279)	(1274)
40th	8411	-2102	-1925	6737	-1339	-1307
	(7)	(19)	(17)	(909)	(1351)	(1348)
50th	9395	-1826	-1736	7415	-741	-693
	(8)	(19)	(19)	(870)	(1487)	(1472)
60th	10555	-1654	-1540	8013	133	118
	(8)	(21)	(20)	(1025)	(1762)	(1703)
70th	11895	-1537	-1296	8673	1336	1274
	(10)	(24)	(23)	(1086)	(2168)	(2163)
80th	13676	-1272	-1010	9477	3615	3394
	(13)	(30)	(29)	(1389)	(3150)	(3190)
90th	16999	-1249	-653	10538	8830	8195
	(20)	(47)	(44)	(2060)	(4458)	(5365)
95th	21234	-1694	-549	11341	14018	13285
	(35)	(86)	(68)	(3095)	(4964)	(9195)

NOTE: N=619,499. The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and whose first birth was a single birth. All the regressions (except where indicated) include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Normalized income is windsorized at zero. QTE regressions use broken first marriage as the first-stage dependent variable. Robust standard errors in parentheses.

Table A. Descriptive Statistics—Mothers with Minor Children Living at Home

	All	Our sample
Demographics		
Age	35.1	31.6
Years of schooling	12.0	12.8
Household income	22,747	23,114
Total own income	4,905	4,458
Weeks worked last year	23.9	22.5
Usual hours worked	21.1	19.9
Marital status		
Currently married, spouse present	0.802	0.891
Currently separated	0.040	0.024
Currently divorced	0.085	0.072
Ever divorced	0.215	0.172
Never married	0.043	0.000
Number of observations	1,610,516	619,499

NOTE: First column includes all women observed in the 1980 Census living with at least one minor child. Second column includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and whose first birth was a single birth. Income is in 1980 dollars.

Figure 1. Mothers' Poverty Rates, by Marital Status

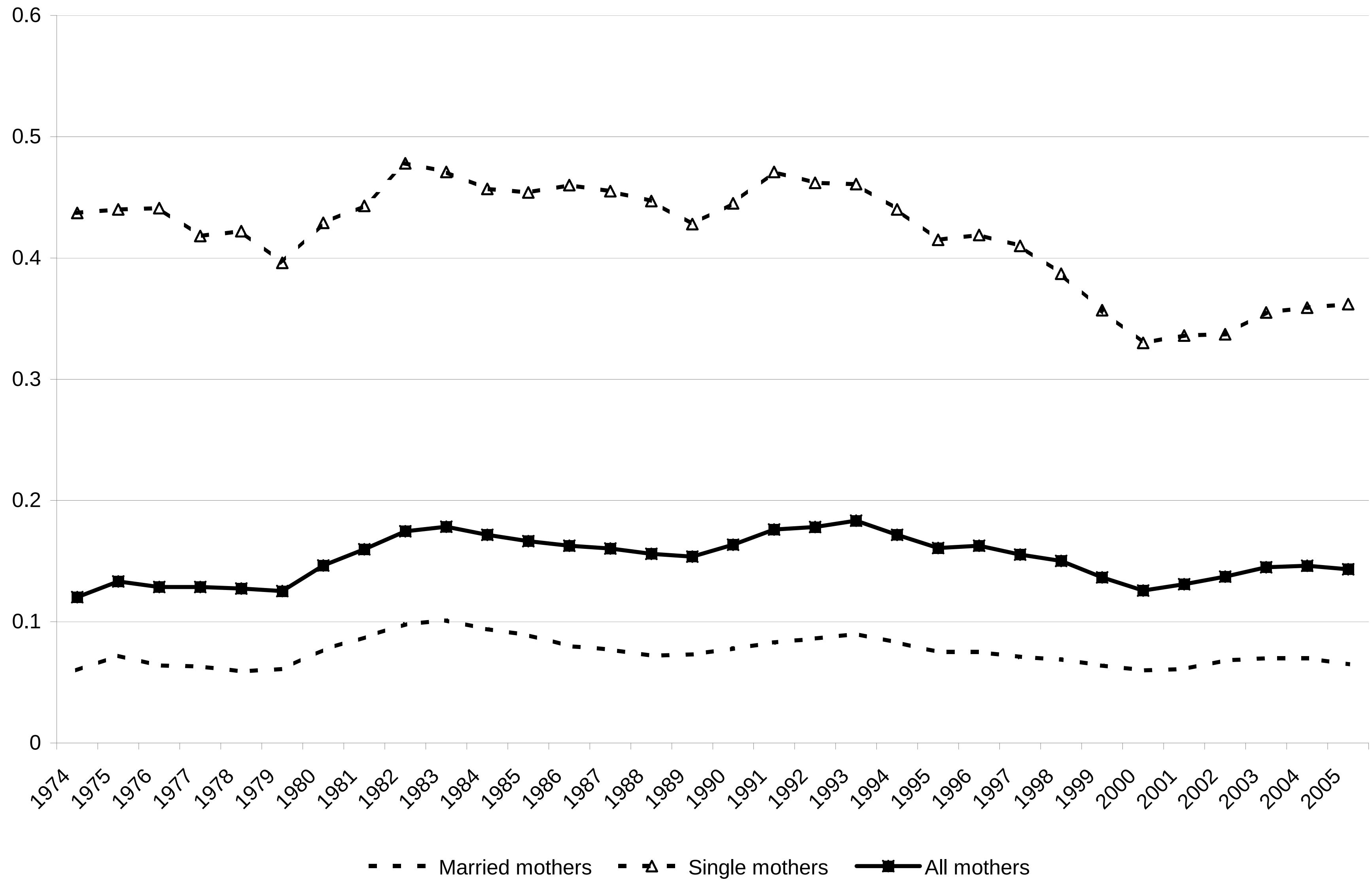

Figure 2. Quantile Distribution of Mothers' Income Sources--Estimates from QR and QTE