The Effect of Marital Breakup on the Income and Poverty of Women with Children

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Abstract

Having a female firstborn child significantly increases the probability that a woman's first marriage breaks up. We exploit this exogenous variation to measure the effect of marital breakup on women's economic outcomes. We find evidence that breakup has little effect on a woman's average household income, but significantly increases the probability that her household will be in the lowest income quartile. While women partially offset the loss of spousal earnings with child support, welfare, combining households, and substantially increasing their labor supply, divorce significantly increases the odds of household poverty on net.

Keywords: divorce, income inequality, labor supply.

JEL classification: J12, J13, J22, I30.

1. Introduction and Motivation

Over the last several decades, divorce has become a much more commonplace event in the United States: In 1950, 2.4 percent of women over age 15 were divorced, while 65.8 percent were married. By 1980, 6.6 percent of women over age 15 were divorced and 58.9 percent were married. And in 2000, those numbers had reached 10.7 percent and 54.1 percent, respectively.

Cross-sectional comparisons suggest strong relationships between marital status and many economic outcomes. For example, single mothers are significantly more likely to work than are married mothers. And they are much more likely to be poor: in 2002, the poverty rate among married women with children was 7.5 percent, while the rate for unmarried women with children was nearly five times as high, at 35.2 percent (U.S. Bureau of the Census 2003). Poverty, in turn, has negative effects on children's health and educational outcomes (e.g. Duncan and Brooks-Gunn 1997).

Current political discussions commonly assume that marriage has causal beneficial 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 current TANF reauthorization bill provides $1.5 billion in funds for "Healthy Marriage Promotion," and many states are exploring their own marriage-promotion initiatives. These initiatives tend to focus on enhancing relationship stability among engaged and married couples.[1] There is also perennial debate over the income tax's marriage penalty, which many argue deters marriage and attendant benefits.

Despite the assumptions made in popular debate, a causal relationship between economic well-being and marriage preservation has not been established. Indeed, there are reasons to suspect that the correlation between marital breakup and poverty overstates the causal effect of divorce. For example, negative household income shocks such as job loss or falling real wages may increase marital tension and thereby increase the likelihood of divorce, so that poverty causes divorce rather than divorce driving poverty. In addition, women with worse earnings opportunities may be less attractive marriage partners, which would mean that women who get divorced have lower income, but not simply because they got divorced. Of course, the possibility also exists that the causal relation is understated.

This paper uses a new instrumental variables approach to separate the causal effects of divorce from its well-known correlations. Using 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. 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 3.2 percent increase from a base likelihood of 20 percentage points. This result, which we will argue implies that sons have a stabilizing effect on the family unit, is consistent with the finding (established in the psychology and sociology literature) that fathers bond more with sons.

When we use child sex as an instrumental variable to measure the effect of divorce on economic outcomes, we find that the marital breakup driven by having a girl does not significantly affect mean household income. Marital breakup does, however, significantly affect the household income distribution; in particular, it dramatically increases the probability that a mother will end up with very low income. While child support, welfare, and an increase in her own earnings (due to a substantially increased labor supply) go some way towards alleviating the loss of her husband's earnings, they often do not prevent poverty status. We also find some evidence that women who divorce may be slightly more likely to end up at the top of the income distribution, possibly due to re-marriage.

Several concerns might reasonably be raised about causal interpretation of the relationships between child sex, marital breakup, and economic outcomes. First, one might be concerned that the observed relationship between child sex and divorce is in fact an artifact of differential custody rates; women are more likely to lose custody of their boys in divorce, leading to divorced women being observed with fewer boys. However, we find highly comparable estimates in CPS supplements that provide detailed fertility histories, which lends evidence that our results are not due to bias in observed household composition. Second, biologists (Trivers and Willard 1973) have argued that child sex is endogenous to resources—that is, that natural selection has favored the ability of parents to have more girls in bad economic times. We find, however, that in our sample there is no difference between the economic resources of mothers of boys and mothers of girls when their first child is born; rather, the difference in income builds up over time, as does the difference in divorce rates. Third, one might worry that sex of the first-born child affects economic outcomes through channels other than divorce. For example, parents may work harder when they have a son. But we find a negligible effect of child sex on economic outcomes in a sample of women with a low exogenous probability of divorce, which provides evidence that child sex is not affecting economic outcomes through a channel other than marital breakup.

After discussing our findings, we consider the potential macro implications of the relationship we identify between divorce and poverty. In recent decades, mothers' poverty rates have failed to decline as much as the overall rate. Can the increase in the proportion of divorced women in the U.S. explain this stagnation? We calculate that, in fact, the poverty rate of women with children today is substantially higher than it would be if divorce had not risen over the post-1980 period.

The paper proceeds as follows. In section 2, we set our findings in the context of previous research. In section 3, we describe the data and the sample we use. In section 4, we describe the estimation strategy for analyzing mean outcomes and describe the results. In section 5, we develop a framework for estimating distributional outcomes and discuss the results. In section 6, we conclude.

2. Divorce and Children's Gender

Previous research has established the need for instrumental variables when examining the effect of marital status on women's outcomes, both in theory (e.g. Becker, Landes, and Michael 1977; Becker 1985) and empirically (e.g. Angrist and Evans 1998). In particular, Gruber (2000) emphasizes the necessity of alternatives to cross-sectional analysis when measuring the effect of divorce on child outcomes.

We instrument for the breakup of a woman's first marriage using the sex of the first-born child. Having a first-born girl could be positively correlated with breakup of the first marriage for one or more reasons involving both the husband and the wife.

First, a man with a daughter may be less interested in sustaining marriage than a man with a son. He may intrinsically value a son more, perceiving him as an "heir." Or he may have more in common with a boy than with a girl, and therefore bond better over time. Second, a woman with a daughter may be less interested in sustaining marriage than a woman with a son. For example, she may believe a son needs his father as a role model, and therefore work harder to retain her husband. Or she may get better companionship from a daughter, and therefore place less weight on her relationship with her husband.

The previous literature supports the idea that fathers tend to bond more with their sons. Aldous, Mulligan and Bjarnason (1998), using data from the US National Surveys of Families and Households, find that fathers spend more time with boys than with girls, and that the gender gap in paternal attention increases as the child ages. This supports the idea of increased bonding over time. Harris, Furstenberg and Marmer (1998) construct a "parental involvement index" and find that fathers' involvement with their adolescent children is higher with sons than with daughters, while mothers' involvement does not differ by child sex. This supports the claim that fathers bond more with sons, but does not lend evidence that mothers get better companionship from daughters. Thus the literature is consistent with the theory that fathers bond more with sons, and that this has a stabilizing effect on the family unit.

Previous work (c.f. Angrist and Evans 1998) considers child sex as an exogenous variable when analyzing U.S. data and after conditioning on race. This seems particularly plausible when considering only births that took place prior to the 1980's, when ultrasound was limited to problematic pregnancies (Campbell 2000). Nevertheless, concern has been raised by some that child sex is endogenous to resource levels. Trivers and Willard (1973) argue that a male is likely to out-reproduce a female if both are in good condition, and vice versa if both are in bad condition. They argue that natural selection may therefore have favored parental ability to adjust sex ratios so that more girls are born in bad times, and they argue that data from mammals support their hypothesis. Our findings, however, are not consistent with the Trivers and Willard hypothesis. We find no relationship between economic circumstances at birth and child sex. Our findings are consistent with the hypothesis that the sex of the oldest child is exogenous.

Dahl and Moretti (2003), in simultaneous research, have examined the relationship between child sex and divorce—that is, our first stage. They too find a significant relationship between having female children and divorce, although our estimates of the effect size are smaller because we address serious selection problems that otherwise overstate the relationship.

In other concurrent research, Bedard and Deschenes (2003) have used the same instrumental variable for divorce: gender of the first child. A recent version of their paper also looks at some of the same economic outcomes. When we use specifications similar to theirs, our results are consistent with their findings: they, like we, find that divorce has no significant negative causal effect on women's mean economic outcomes and that the cross-sectional relationship between divorce and household income at the mean is due to selection.

However, when we look beyond the mean to consider the way in which marital breakup affects the entire income distribution of women with children, we arrive at a different conclusion. Namely, we find that marital breakup has significant negative consequences for many women; in particular, it increases the probability that a woman lives in poverty.

We also find some evidence that women who remarry may actually become "better off" in terms of income after the end of their first marriages. This is consistent with an earlier literature that does not attempt to establish causality or deal with selection into remarriage: Mueller and Pope (1980) and Jacobs and Furstenberg (1986) find that women who remarry do better in terms of their husbands' education level and SES score. Duncan and Hoffman (1985) find that five years after their divorce, women who remarried gained in terms of family income.[2]

We are unaware of any other work that looks at the causal relationship between divorce and inequality, although both have increased substantially over the past three decades. Much of the recent literature on the causes of 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 decline of the traditional family unit as an institution may have contributed to the rise in income inequality.

3. Data

We use data on women living with minor children from the 5 percent 1980 Census file, which allow us sufficient power to identify the effect of sex of the first-born child on marital breakup.[3] We limit our sample to white women who are living with all of their children, whose oldest 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. In particular it is important to consider only women who live with all their children, since by doing so we avoid any spurious correlation between child sex and household structure due to differential maternal custody rates of boys and girls. For further discussion of the sample construction, see the appendix.

In addition to estimating our model on the full sample, we look specifically at two subsamples that we use in specification checks: those who are at high risk of having ever divorced and those at low risk of having ever divorced. We create an index of exogenous risk for divorce (see Appendix). The high-risk subsample includes those whose predicted risk is in the top quartile; the low-risk subsample includes those whose predicted risk is in the bottom quartile. These subsamples allow us to test the validity of our IV strategy. We hypothesize that the effect of child sex on economic outcomes should be stronger for the group which has a higher divorce rate—because the marital breakup mechanism through which child sex affects income is stronger. Similarly, child sex should have a weaker effect on economic outcomes in the low-divorce subsample.

Table 1 shows summary statistics for our full sample and subsamples 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. Our sample is younger than average, consistent with the requirement that a woman's oldest child is under 17. The women in our sample are 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.

The high-predicted divorce subsample is much older on average than the low-predicted divorce subsample, consistent with more years of exposure to risk of divorce. Women in the high-predicted divorce subsample have less education—probably due to cohort effects as well as the fact that education enters negatively into the divorce prediction—but higher own and household income, probably due to lifecycle effects.

4. Estimation of the Effect of Marital Breakup on Average Economic Outcomes

Estimation Framework

We begin our investigation of the causal effects of divorce by considering a simple econometric model. In this model, income is affected by the breakup of the first marriage, which is treated as a classic endogenous regressor:

$$(1) \qquad D = \alpha_1 Z + X\alpha_2 + U\alpha_3 + u$$

$$(2) \qquad Y = \beta_1 D + X\beta_2 + U\beta_3 + \varepsilon$$

The right-hand-side variable of interest in equation (2), D, is a dummy for the breakup of the first marriage. We define 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 define as having her first marriage broken any woman who: 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.[4]

Our outcomes, Y, are total others' income, defined as total household income less total own income; household income; a measure of household poverty; and in some specifications hours worked. The first outcome, others' income, measures the direct effect on a woman of losing her husband as a source of income (to the extent that the husband is not replaced by other wage earners). The second outcome, total household income, also includes the indirect effects of divorce on income: these include transfers from the ex-husband in the form of alimony and child support[5] and from the state in the form of cash assistance as well as income generated by her own labor supply response.[6]

Our controls, denoted by X, are a vector of pre-determined demographic variables including age, age squared, age at first birth and a dummy for high-school dropouts.[7] While our OLS estimates of the relationship between marital breakup and income may be sensitive to their inclusion, our IV specifications are robust to controls (results without controls are available upon request). 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, Z represents our instrumental variable, the sex of the firstborn child (male=0, female=1), which (we will show) exogenously affects the likelihood of divorce.

As discussed above, there are two major problems with estimating equation (2) with OLS. The first, correlation of U and D, can be seen as omitted variables bias or a selection problem, and it induces a bias of indeterminate direction. Women with worse earnings opportunities may be less attractive marriage partners: this would imply negative selection into marital breakup and upward bias in the estimated cost of marital breakup for women. On the other hand, women with worse earnings opportunities may be more interested in sustaining marriage. Or a stronger preference for household work relative to market work may cause a woman to work harder at staying married. In this case, the women who stay married would have supplied fewer hours of labor had their marriage broken up, and their earnings would have been lower.

The second major problem with OLS estimates of equation (2) is reverse causality. Conditional on a woman's observed and unobserved characteristics, a negative income shock to the household may increase marital tension and thereby raise the likelihood of divorce.

To address concerns both about omitted variable bias and about reverse causality, we estimate equation (2) using two-stage least squares, using the sex of the oldest child (Z) as an instrument for whether the first marriage is broken (D). Angrist and Imbens (1994) show that in the absence of covariates and given the standard two-stage least squares assumptions,[8] the IV approach identifies the local average treatment effect:

$$(3) \qquad \beta_{1,IV} = \frac{Cov(D,Y)}{Var(D)} = \frac{E[Y \mid Z=1] - E[Y \mid Z=0]}{E[D \mid Z=1] - E[D \mid Z=0]} = E[Y_1 - Y_0 \mid D_1 > D_0]$$

Here Y_1 and Y_0 denote the income (or other dependent variable) for women whose first marriage is broken and intact, respectively. D_1 is an indicator for whether a woman would divorce if her first child were a girl; D_0 is an indicator for whether she would divorce if her first child were a boy. In practice, of course, we only observe the indicator for the child sex that is realized.

β_{1IV} therefore measures the change in income due to divorce for women whose first marriage breaks up if they have a girl and remains intact if they have a boy. That is, two-stage least squares estimates the average responsiveness only of that part of the population for whom the treatment is equal to the instrument ($D_1 > D_0$), the group known as "compliers." As one might imagine, this group is typically not a random subsample of the population, and in fact we will show below that those who comply with sex of the first-born child tend to be of low socio-economic status.

<u>First Stage Results: Child Sex and Marital Breakup</u>

The first stage results for the entire population as well as for the full sample and various subsamples are presented in Table 2. When equation (1) is estimated for all women living with minor children, the coefficient on child sex is 1.13 percent. We believe that this figure overstates the actual effect of child sex on marital breakup, because endogeneity in the sex of the oldest child residing with a woman will create a spurious correlation between living with an oldest girl and being divorced.[9] That is, since women are more likely to end up with custody of girls than boys in the event of divorce, living with a girl is cross-sectionally correlated with being divorced above and beyond the causal effect of having a girl on marital breakup.

The second column gives the estimated relationship for our sample, which is constructed in order to close the custody channel, so that the relationship between child sex and marital breakup is causal. We find that the average effect of the first child being a girl on the breakup of the first marriage is about 0.63 percent, which is substantially smaller than the full-population estimate, but still highly significant. Furthermore, as discussed below, we replicate our result using data from the Current Population Survey, which identifies the sex of a woman's actual firstborn child (and hence requires no sample restrictions). The CPS results are highly similar to those from our Census sample.

The effect increases with age, suggesting, sensibly, that the sex of the first child affects the hazard rate of divorce and separation. This can also be seen in Figures 1 and 2, each presenting the results from 17 separate regressions by the age (0 to 16) of the oldest child. We also find that the effect size varies by the woman's education—high-school dropouts exhibit the strongest effect, and the effect size diminishes as education increases. This suggests that our instrument primarily provides us insight on the effects of marital breakup particularly for low-SES women.

Finally, we find a larger and more significant effect on those at higher exogenous risk of divorce: the quarter of our sample with the highest predicted risk of divorce has a first stage of 1.13 percent (highly significant), and the group that is least likely to divorce has a first stage of 0.35 percent (marginally significant). The differential effect by underlying probability of divorce lends evidence that the instrument is acting in the way we would anticipate.

One further implication of the figures should also be noted: a comparison between Figures 1 and 2 reveals that the effects of sex of the oldest child on the breakup of the first marriage are bigger than those on being currently divorced. Much of this difference comes from the fact that some of the women who divorced due to the instrument have since remarried. In addition, some of those who have not re-married have instead moved in with other adults (such as their parents). We find that having an oldest girl increases the probability that a woman is the head of household by only about half as much as the increase in probability that her first marriage is broken (result not shown).[10] When interpreting the second-stage results below, therefore, it is important to keep in mind that we are measuring the effect of having ever experienced a broken marriage, not the effect of being currently divorced or of residing in a mother-only household.

Specification Checks

While the lack of over-identification implies that we cannot formally test the validity of our instrument, we do bring to bear some evidence that the instrument is not correlated with other omitted variables. First, we test the Trivers and Willard hypothesis that adverse conditions cause women to give birth to a higher proportion of girls. We have examined the income of mothers who were never married, and found that the income distribution in this group was very similar whether the oldest was a girl or a boy. We have also examined the income of women who were married when their first child was born and whose first marriage was still intact, and similarly found that the income distribution in this group was virtually the same whether the oldest was a girl or a boy. (Results not shown.)

This is still not a conclusive refutation of the endogenous sex hypothesis for our sample: the first group, those who were never married, are not included in our sample (we limit to those whose first birth occurred after their first marriage); the second group includes women who remain married after having a girl (i.e. those who do not respond to our instrument, and thus for whom we cannot measure the causal effect of divorce). But we have also found that the reduced-form effect of the instrument on household income and on others' income is small and statistically insignificant shortly after the child is born. Rather, the effect of the instrument increases with time.[11] This suggests that child sex is not the result of resources, but rather that later resources are the result of child sex, through its effect on marital breakup—note that the effect of sex on breakup similarly builds over time (see Figures 1 and 2).

If child sex is not the result of resources, we can interpret as causal the reduced-form relationship between having an oldest girl and economic outcomes. But omitted variables might mean that child sex affects outcomes through paths other than divorce, so that our IV estimate of the effect of divorce on outcomes is overstated. For example, parents may work harder in the labor market when they have a boy. Alternatively, having a girl may, even in the absence of divorce, cause increased marital conflict, which could in turn negatively impact economic outcomes. In either case, having a girl would be negatively correlated with women's outcomes, but not solely through marital breakup, and using it as an instrument would cause an overstatement of the cost of divorce. We again cannot refute either of these possibilities directly. However, when we looked at the sample of women whose marriages remain intact, the income distribution of mothers did not substantially differ by the sex of the oldest child.

We might also overestimate the cost of divorce if fathers pay differential child support by the sex of the oldest child. It is possible that fathers of boys transfer more money to the mother of their child, even if they are no longer married to her. In this case, again, having a girl would be negatively correlated with women's outcomes, but not solely through marital breakup. To check this possibility, we examined the Census "other income" variable (which includes child support and alimony) for women who are currently divorced and for women who are ever divorced. We found that other income did not differ by the sex of the oldest child (results not shown).[12]

A final caveat remains for using the sex of the oldest child as an instrument for marital breakup: having a son as one's oldest child may increase the likelihood of re-marriage after the first marriage was broken. Note, again, that this would not invalidate the causal interpretation of the reduced-form estimates, but could cause us to overstate the direct effects of divorce. Again, we cannot rule out the existence of such a causal effect. But when we regress a dummy for being currently married on having an oldest girl for ever-divorced women in our sample, the coefficient is indeed negative, but is not statistically significant (t-statistic=0.67).

As a final check on the validity of the relationship between child sex and marital breakup, we replicate our Census estimates using CPS data. We do so to address potential concern about the reliability of our Census sample—that despite our best effort to construct a set of women for whom error in observed oldest child sex is uncorrelated with marital status, some error remains and is contaminating our estimate. We address this concern by using CPS fertility supplements. Because they record a woman's fertility history regardless of whether her children are still in the household, we do not have to worry about bias in the sex of the children still in the household.

Four June CPS supplements (1980, 1985, 1990, and 1995) record the sex of a woman's first-born child. In this sample, we need only restrict to women who are white, had their first birth after their first marriage, and had a single first birth—we need not make any of the Census restrictions that were geared towards assuring that we accurately observe the sex of the first child.[13] When we pool these supplements (N=107,519), we again find that having an oldest girl significantly increases the probability of marital breakup. In fact, we find that having an oldest girl causes a 0.68 percent increase in the probability that the first marriage breaks up (t-statistic=2.31). We consider the similarity of the point estimates from both samples (the Census estimate is 0.63 percent) an encouraging sign that our results reflect a real and significant effect of the sex of the first child on the probability of marital breakup.

The effect of marital breakup on mean outcomes

Cross-sectional (OLS) regressions of income and labor supply on marital breakup (Table 3), which admit no causal interpretation, show that breakup of the first marriage is correlated with large losses in income and large increases in labor supply. Both of these relationships confirm the conventional view that women whose first marriages end are significantly worse off than women whose first marriages remain intact.

Two-stage least squares estimates, on the other hand, suggest that the mean effect of marital breakup on material well-being is quite different from the cross-sectional results. The two-stage estimate of the effect of divorce on others' income is negative but insignificant, though still within two standard deviations of the OLS estimates, as is the two-stage estimate for log household income.

The two-stage estimate of the effect of divorce on household income level, however, is significantly more positive than the OLS estimate—while the coefficient is not statistically different from zero, it is more than two standard deviations from the negative OLS estimate. The same is true for the high predicted-divorce subsample. For the low predicted-divorce subsample our first stage is barely significant, so zero effects or effects equal to the cross-sectional effect cannot be ruled out.

Taken together, these results imply 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. The lack of significant effect on others income is probably due to the fact that most women compensate for the loss of their husbands' income by re-marriage or by moving in with other adults. Moreover, we find a very large increase in mean hours worked (about 1,000 hours a year), which is more than double the cross-sectional effect and clearly helps offset the loss of husband income in the measure of household income.

As the labor supply results indicate, the stability of mean income should not be taken to imply that there is no mean welfare loss from divorce. The increase in work without a significant increase in income implies welfare loss for women, since they experience a mean decrease in leisure without an expansion of consumption possibilities. In addition, it implies potential adverse consequences for children, since they too have no obvious gain in consumption but likely experience a decrease in time with their mother (and probably with their father as well).

5. Estimation of the Effect of Marital Breakup on the Income Distribution

<u>Estimation Framework</u>

The difference in sign between the causal effect of divorce on mean log income and that on mean level of income leads us to investigate the possibility that there are important effects of marital breakup on the income distribution, particularly at the bottom, that aren't evident at the mean. In fact, it makes intuitive sense that the effect of divorce on the income distribution would be to fatten the lower tail, rather than to shift the entire distribution uniformly downward. 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 other relatives, or have high earning potential, many may end up as well off financially as before (or even better off)—resulting in little effect of divorce on the mean of the distribution. And yet a subset of women who cannot recover from the loss would experience a much greater than average effect of divorce, falling near the bottom of the distribution.

To address this possibility, we consider more flexible specifications in which we estimate the effect of marital breakup on the probability that a woman's income is below various thresholds. This enables us to identify the marginal effect of marital breakup on the CDF. In addition, we obtain estimators of the CDF itself for those whose first marriage is broken and for those whose first marriage is intact. Doing so allows us to evaluate the size of the effect of divorce on the income distribution.

Recall that the local average treatment effect formula in equation (3) gave the causal effect of treatment on the compliers. Similarly, Abadie (2002, 2003) demonstrates that in absence of covariates and with the same assumptions as the standard two-stage least squares model:[14]

$$(4) \qquad E[Y_0 \mid D_1 > D_0] = \frac{E[(1-D)Y \mid Z=1] - E[(1-D)Y \mid Z=0]}{E[(1-D) \mid Z=1] - E[(1-D) \mid Z=0]}$$

We can similarly estimate the following equation using two-stage least squares with controls, using the sex of the oldest child as an instrument:

$$(5) \qquad (1-D)Y = \gamma(1-D) + X\delta + \mu$$

This strategy gives a consistent estimator:

$$(6)\qquad \hat{\gamma}_{IV} \rightarrow \gamma = E[Y_0 \mid X; D_1 > D_0]$$

That is, γ gives the expected value of Y for compliers who have a boy (and whose marriage therefore remains intact). When we apply this method to the case where Y is the CDF of the income distribution we can effectively trace out the CDF for compliers who have boys. Similarly, we could estimate the CDF for compliers who have girls.[15] But a standard two-stage least squares estimate of the effect of marital breakup on an indicator for a given income level gives the difference between the two CDFs. Thus, by summing the standard coefficient and the estimate of equation (6), we can likewise trace out the CDF for compliers who have girls. Finally, we also estimate equation (6) using OLS, and compare it to the instrumental variables estimate.

The effect of marital breakup on the distribution of outcomes

In each of Tables 4, 5, and 6, the columns of results labeled "First marriage intact" report the CDF for those who remain married, estimated using the method described above. (Recall that in the two-stage least squares regressions, the CDFs 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 columns labeled "Difference: broken – intact" report the estimated difference in CDFs between the compliers who have divorced and those who stay married, which is the standard coefficient on breakup in the OLS or two-stage least squares estimation.[16]

OLS estimates of the difference in distribution of income by marital breakup (first two columns in Tables 4 and 5) tell a familiar story. In cross-section, those with broken first marriages have income distributions —both household and others'— that are first-order stochastically dominated by those of women with intact first marriages. That is, divorce is correlated with a uniform shift downward in income at every point in the income distribution.

The two-stage least squares estimates in Tables 4 and 5 give a more nuanced picture. We do in fact find a large effect of marital breakup on the probability of being at the bottom of the income distribution: almost half of compliers who divorce have zero earnings of others in the household, compared to virtually none of the compliers who stay married. In addition, women whose first marriage is broken are 42 percentage points more likely (ten times as likely) to have less than $5000 in others' income, and 23 percentage points more likely (60 percent more likely) to have less than $10,000. While legal transfers (which include child support and mean-tested transfers) reduce the number of compliers with no income, over a quarter of the divorced compliers have no unearned income even after accounting for transfers—again compared to virtually none of the compliers who stay married (results not shown).

The analysis of household income similarly shows a large effect of marital breakup on the density at the bottom of the income distribution. While neither type of complier has a significant likelihood of zero household income, 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.[17] Results follow a similar pattern for the high predicted-divorce subsample. For the low predicted-divorce subsample, the effects are not significant.

Interestingly, the two-stage estimates also show that those who divorce due to the instrument (in both the full sample and the high predicted-divorce subsample) are somewhat *more* likely to have income near the top of the distribution,[18] although the differences are not statistically significant. The reversal in the sign of the difference occurs at (in the case of others' income) or below (in the case of household income) the mean of the distribution, explaining why models for the mean find little effect of marital breakup on income. Previous literature (cf. Mueller and Pope 1980, Jacobs and Furstenberg 1986) has found that when divorced women remarry, their second husband is typically more educated and has a higher occupational SES score (although at least some of this is due to life-cycle effects). The reversal in sign is more substantial for total household income than for others' income, probably because that also includes top-end variation in women's earnings as well as sources of other income (such as alimony and child support).

The results discussed thus far neglect one important aspect of divorce, however: it reduces family size. Thus even though income decreases at the bottom with divorce, women may not necessarily end up worse off; on the other hand, if income gains at the top come primarily through remarriage, the effect may be neutralized by increased family size. To estimate the effect of divorce on the ratio of income to needs, we divide each woman's total household income by the poverty line for a household of that size.

As shown in Table 6, we find that changes in family size do not mitigate our results. The OLS estimates still indicate that marital breakup decreases normalized household income at all levels. The two-stage least squares results indicate that virtually none of those 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. Compliers whose first marriage ended are, however, significantly more likely to be above 400 percent of poverty than are compliers whose first marriage remains intact.

6. Discussion

Our results suggest 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: divorce increases the percent of women at the bottom—and perhaps at the top—tail of the income distribution. In net, divorce causally increases poverty, and perhaps inequality more generally, for women with children.

How much lower would the poverty rate of women with children be today if the divorce rate had stayed at its 1980 level? Within our sample, 17.2 percent of women were ever divorced; for a similar sample from the 1995 CPS (the latest year for which representative statistics are available), the rate is 28.6 percent. Our two-stage least squares point estimate suggests that the probability of having income below the poverty line is 12.4 times as high when a woman's first marriage has ended; however, our estimated effects are not statistically different from the cross-sectional OLS estimate that the ratio of poverty rates for these groups is 3.6.

We make the conservative assumption that the true effect of divorce on poverty equals the cross-sectional effect for the purposes of calibration. In 1995, the poverty rate for women with our sample characteristics who are still in their first marriage is 8.7 percent. Thus if the divorce rate had stayed at its 1980 level, the overall poverty rate in this sample would be .172*(3.6*8.7) +.828*8.7, or 12.6 percent. Because the divorce rate rose to 28.6 percent (an increase of 66 percent from the 1980 base), the poverty rate became .286*(3.6*8.7) +.714*8.7, or 15.5 percent. Thus the increase in divorce may potentially have caused an increase of 2.9 points, or 20.5 percent, in the poverty rate for women with our sample characteristics. If we assume that the effect holds outside of those with our sample characteristics, we can conclude that roughly 2.7 million more women and children were in poverty in 1995 than would have been if the divorce rate had remained at its 1980 level.

We conclude by noting that our findings do not have clear policy implications. While some might interpret them as evidence for the importance of traditional family structures, others may conclude that they emphasize the need for more generous welfare policies.

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 of first marriage. This information permits us to identify the sex of the first-born child for most women, although not for women whose oldest 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 residing with a woman, whereas our true characteristic of interest is the sex of the firstborn child. It is imperative 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 may be differentially likely to end up in the custody of their fathers in the event of marital breakup. In fact, as Table A, panel 1, shows, girls are significantly less likely to be living with their father and not their mother. 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 oldest *observed* child being a girl.

To address this issue, we exclude 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 are living with all of their children reduces but does not eliminate the threat that differential custody rates could bias our result—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 find, however, that mothers living with all their children and mothers in the overall population are equally likely to be observed with a girl as the oldest child, which suggests that sex of the oldest child is not a major determinant of living with all of one's children (See Table A, panel 2).

Second, we limit our sample to women whose first child was 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. 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. But 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. But this should create only classical measurement error in our first stage estimation.

Third, we look only at mothers whose oldest child is a minor, since those who still live with their adult children may be a select group. Further, since girls are differentially likely to leave home early (at ages 17 and 18), we restrict to mothers whose oldest child is under age 17. Fourth, we limit our sample to white women because black women's childbearing and marital decisions may be quite different, and fully modeling the differences would greatly complicate the analysis. Finally, we leave out women 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.

Selection into our restricted sample might be cause for concern, just as selection into the labor market is cause for concern when measuring labor outcomes. To test for such a problem, we generated the predicted probability that a woman was included in our sample based on age and birthplace dummies. Then we re-ran our two-stage estimates, treating a woman's predicted probability of inclusion as an endogenous regressor. Our first- and second-stage estimates were quite stable with and without this variable.

To create our high- and low-predicted divorce subsamples, we create an exogenous risk index (Table A, panel 3) by predicting "ever divorced" using age, age squared, age at first birth, and dummies for place of birth and for high school dropout (recall that our sample includes only women whose first birth occurs after age 19, so the decision on whether to graduate from high school can be treated as pre-determined).

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[1] "Bush plans $1.5 billion drive for promotion of marriage." New York Times news article, 1/14/04; 2003 H.R. 4: http://frwebgate.access.gpo.gov/cgi-bin/getdoc.cgi?dbname=108_cong_bills&docid=f:h4rfs.txt.pdf.

[2] Both Jacobs and Furstenberg and Duncan and Hoffman argue that this is largely due to life cycle effects, so that *if* the man these women divorced is similar to the average person in his age group, he would have made similar gains from the time of his divorce until the time we observe the woman's second husband. We suspect, however, that in our case the men who divorce are negatively selected, so it is plausible that women who re-married did somewhat improve their economic situation.

[3] While earlier censuses have these measures, in previous decades the divorce rate was very low. Subsequent censuses, on the other hand, don't have all of the necessary measures. We did, however, try similar specifications using 1990 data and found similar

patterns to those reported here; this finding lends evidence that cohort-specific effects are not driving the relationship between firstborn sex and divorce.

[4] 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.

It is not possible to systematically remove widows from the sample, since the data do not allow us to identify those whose first marriage ended in death among those who have had multiple marriages. In any event, since widowhood is endogenous to both socioeconomic status and marital duration, it is probably not desirable to exclude widows.

[5] 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."

[6] We have also looked at various intermediate outcomes, such as the level of alimony and child support, the level of welfare, the woman's own earnings, and the woman's hours. All of the analyses gave results highly consistent with the results presented here, and are available from the authors upon request.

[7] Since we look only at women who gave birth to their first child after age 19, 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.

[8] The assumptions are: conditional independence of Z, exclusion of the instrument, existence of a first stage and monotonicity – see Abadie (2002). The result can be generalized for the case with covariates.

[9] In fact, a regression similar to those in Table 2 for a sample of women with all of our restrictions except that she must reside with all of her children gives an estimate of .010. This result is much closer to the estimate in column 1 than in column 2, suggesting that the factor driving the difference between the estimates in the full population and in our sample is the endogeneity of child residence.

[10] Note that this will not capture all avenues of combining households, since a woman may share a household and be identified as its head.

[11] We also regressed the sex of the oldest child on household income and controls for mothers whose children were aged 0-1 and found no significant effect of household income on child sex.

[12] Dahl and Moretti (2003) find that divorced mothers of multiple children who are all girls are marginally significantly less likely to receive child support than those whose children are all boys. They, however, use a much less restricted sample, and use the gender of all children. Even within their sample and methodology, when they examine all family sizes they do not find a significant effect.

[13] Compared to the census data, the CPS has drawbacks as well as advantages. On the one hand, the CPS allows us to look at older women, whose oldest children left their house after their full impact on marital stability had been realized, thus increasing the potential causal effect. On the other hand, older women are less likely to have experienced divorce (due to cohort effects) and widowhood may be higher, attenuating the estimated effect of

child sex on marital breakup. Finally, in order to get a large enough sample, we needed to look at more recent data, since the CPS recorded women's fertility history only in the four samples we used.

[14] Proof of equation (4): $Z = i \Rightarrow D = D_Z \Rightarrow Y = Y_Z$ for $i = 0,1$, so the numerator is:

$$E[(1-D)Y \mid Z=1] - E[(1-D)Y \mid Z=0] = E[(1-D_1)Y_1 - (1-D_0)Y_0]$$

By the monotonicity assumption this equals:

$$E[(1-D_1)Y_1 - (1-D_0)Y_0 \mid D_1 > D_0]P[D_1 > D_0] = E[Y_0 \mid D_1 > D_0]P[D_1 > D_0]$$

Dividing by the denominator yields: $E[Y_0 \mid D_1 > D_0]$.

Since X is discrete, this proof can be extended to the case where we add controls.

[15] By estimating the equation: $DY = \theta D + X\rho + \eta$ we can get

$\hat{\theta}_{IV} \to \theta = E[Y_1 \mid X; D_1 > D_0]$. The proof is similar.

[16] Note that the estimation of linear probability models results in some estimates that are slightly outside the [0,1] interval, though always within two standard errors of this interval. Despite this drawback, we prefer to follow this methodology because of its transparency.

[17] 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, a separate analysis is not included here.

[18] We have developed a simple model of the marriage and remarriage markets that explains this result as well as our other key findings. It is available upon request from the authors.

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Table 1. Descriptive Statistics—Mothers with Minor Children Living at Home

Variable	All women with children	Samples		
			Predicted-divorce index	
		Full sample	High	Low
Demographics				
Age	35.1	31.6	36.0	25.2
Years of schooling	12.0	12.8	11.5	13.0
Household income	22,747	23,114	23,598	19,549
Total own income	4,905	4,458	4,625	3,805
Weeks worked last year	23.9	22.5	23.1	21.4
Usual hours worked	21.1	19.9	19.8	21.2
Marital status				
Currently married, spouse present	0.802	0.891	0.869	0.913
Currently separated	0.040	0.024	0.028	0.025
Currently divorced	0.085	0.072	0.084	0.052
Ever divorced	0.215	0.172	0.202	0.118
Never married	0.043	0.000	0.000	0.000
Number of observations	1,610,516	619,499	154,041	155,355

NOTE: The full sample includes white women who are living with all of their children, whose oldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. The high predicted-divorce subsample includes only women in the top quartile of the predicted divorce distribution; the low predicted-divorce subsample includes only those in the bottom quartile.

Table 2. OLS Estimates of Marital Breakup on Child Sex

	All women with children	Samples					
			Age of oldest child			Predicted-divorce index	
		Full sample	0-5	6-11	12-16	High	Low
All education levels	0.0113	0.0063	0.0038	0.0057	0.0106	0.0113	0.0036
	(0.0007)	(0.0010)	(0.0015)	(0.0018)	(0.0021)	(0.0022)	(0.0018)
Observations	1,605,339	619,499	236,522	221,937	161,040	154,041	155,355
High school dropouts		0.0149	0.0106	0.0090	0.0264		
		(0.0033)	(0.0054)	(0.0057)	(0.0059)		
Observations		73,426	25,111	24,849	23,466		
High school graduates		0.0057	0.0041	0.0053	0.0084		
		(0.0015)	(0.0024)	(0.0027)	(0.0030)		
Observations		277,155	100,601	99,816	76,738		
At least some college		0.0045	0.0023	0.0052	0.0075		
		(0.0015)	(0.0020)	(0.0026)	(0.0034)		
Observations		268,918	110,810	97,272	60,836		

NOTE: The sample includes white women who are living with all of their children, whose oldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. The table reports estimates of the coefficient on Oldest child is a girl in equation (1) in the text. Other covariates are: Age, Age squared, Age at first birth and High school dropouts. Robust standard errors in brackets

Table 3. OLS and 2SLS Estimates of Economic Outcomes on Broken First Marriage

	Full Sample		Predicted-divorce index			
			High		Low	
Dependent Variable	OLS	2SLS	OLS	2SLS	OLS	2SLS
Ln(Household income)	-0.501	-0.230	-0.547	-0.867	-0.439	1.514
	(0.005)	(0.437)	(0.009)	(0.542)	(0.011)	(1.693)
Others' income	-9241	-1041	-10561	-8450	-5014	8605
	(42.57)	(5178)	(78.68)	(5939)	(87.82)	(15369)
Household income	-5577	6548	-6721	-981	-2649	9147
	(43.41)	(5570)	(80.57)	(6224)	(89.76)	(15817)
Hours worked last year	420	1053	347	937	369	951
	(2.85)	(360)	(5.39)	(408)	(6.31)	(1180)
Number of observations	619,499		154,041		155,355	

NOTE: The full sample includes white women who are living with all of their children, whose oldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. High predicted-divorce subsample includes only women in the top quartile of the predicted divorce distribution; the low predicted-divorce subsample includes only those in the bottom quartile. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. There were 4,766 observations, or about 0.77% of of our observations, for which household income was zero or negative. For those observations we set Ln(Household income) to zero.

Table 4. Estimated CDF of Others' Income

	OLS		2SLS					
					Predicted divorce index			
	Full Sample		Full Sample		High		Low	
Dependent Variable	First marriage intact	Difference: broken-intact	First marriage intact	Difference: broken-intact	First marriage intact	Difference: broken-intact	First marriage intact	Difference: broken-intact
Others' income ≤$0	0.0107	0.3636	-0.0215	0.4863	0.0076	0.6922	-0.0952	0.5042
	(0.0003)	(0.0008)	(0.0371)	(0.0979)	(0.0452)	(0.1326)	(0.1322)	(0.2975)
≤$5000	0.0507	0.3928	0.0378	0.4177	0.0800	0.5249	0.1587	0.2041
	(0.0006)	(0.0009)	(0.0785)	(0.1202)	(0.0930)	(0.1457)	(0.3061)	(0.4067)
≤$10000	0.1458	0.3850	0.2770	0.2284	0.2932	0.2727	0.8545	-0.4356
	(0.0010)	(0.0012)	(0.1270)	(0.1567)	(0.1414)	(0.1785)	(0.6081)	(0.6870)
≤$20000	0.5459	0.2226	0.8751	-0.0416	0.5735	0.1968	1.0290	-0.0898
	(0.0014)	(0.0015)	(0.1814)	(0.1937)	(0.1866)	(0.2054)	(0.5957)	(0.6208)
≤$30000	0.8412	0.0754	1.0369	-0.0821	0.9996	-0.0195	0.8886	0.1992
	(0.0010)	(0.0011)	(0.1323)	(0.1399)	(0.1554)	(0.1629)	(0.3167)	(0.3620)
≤$40000	0.9303	0.0322	1.0368	-0.1144	1.0102	-0.0245	0.9698	-0.0843
	(0.0007)	(0.0008)	(0.0920)	(0.0994)	(0.1128)	(0.1183)	(0.2014)	(0.2268)
≤$50000	0.9610	0.0183	1.0491	-0.1084	0.9618	0.0289	1.1175	-0.2703
	(0.0005)	(0.0006)	(0.0703)	(0.0763)	(0.0867)	(0.0912)	(0.1587)	(0.2082)
Number of observations	619,499		619,499		154,041		155,355	

NOTE: The sample includes white women who are living with all of their children, whose oldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Robust standard errors in parentheses.

Table 5. Estimated CDF of Household Income

	OLS		2SLS					
					Predicted divorce index			
	Full Sample		Full Sample		High		Low	
Dependent Variable	First marriage intact	Difference: broken-intact	First marriage intact	Difference: broken-intact	First marriage intact	Difference: broken-intact	First marriage intact	Difference: broken-intact
Household income ≤$0	0.0057	0.0109	-0.0086	-0.0040	-0.0020	0.0310	0.1364	-0.1242
	(0.0002)	(0.0003)	(0.0268)	(0.0355)	-0.0339	(0.0454)	-0.1075	(0.1307)
≤$5000	0.0294	0.0982	-0.0155	0.1900	0.0082	0.3086	0.2163	-0.0367
	(0.0005)	(0.0007)	(0.0603)	(0.0865)	-0.0739	(0.1136)	-0.2438	(0.3082)
≤$10000	0.0921	0.1999	0.1502	0.2875	0.1941	0.3994	0.6388	-0.2840
	(0.0008)	(0.0010)	(0.1024)	(0.1312)	-0.1194	(0.1591)	-0.4912	(0.5475)
≤$20000	0.3994	0.2153	0.7760	-0.2547	0.5994	-0.0094	1.7693	-0.8853
	(0.0013)	(0.0015)	(0.1806)	(0.2048)	-0.1821	(0.2086)	-0.8696	(0.8329)
≤$30000	0.7496	0.0788	1.1079	-0.2360	0.8813	-0.0610	0.8815	0.2333
	(0.0012)	(0.0013)	(0.1621)	(0.1729)	-0.1740	(0.1897)	-0.4203	(0.4759)
≤$40000	0.8981	0.0277	1.0810	-0.1841	1.1274	-0.1578	0.8413	0.1970
	(0.0008)	(0.0009)	(0.1111)	(0.1217)	-0.1381	(0.1457)	-0.2578	(0.2987)
≤$50000	0.9300	0.0191	1.0963	-0.1790	1.0591	-0.0833	0.9496	-0.0481
	(0.0007)	(0.0008)	(0.0828)	(0.0917)	-0.1014	(0.1081)	-0.1731	(0.1935)
Number of observations	619,499		619,499		154,041		155,355	

NOTE: The sample includes white women who are living with all of their children, whose oldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Robust standard errors in parentheses.

Table 6. Estimated CDF of Income as a Percent of Poverty by Marital Outcome, for Full Sample

Dependent Variable		OLS		2SLS	
		First marriage intact	Difference: broken-intact	First marriage intact	Difference: broken-intact
Percentage of the poverty threshold	≤50	0.0228	0.0558	-0.0565	0.1489
		(0.0004)	(0.0006)	(0.0539)	(0.0733)
	≤100	0.0565	0.1322	0.0041	0.2407
		(0.0006)	(0.0008)	(0.0806)	(0.1081)
	≤200	0.2224	0.2069	0.2958	0.2487
		(0.0011)	(0.0013)	(0.1449)	(0.1684)
	≤300	0.5051	0.1550	0.8077	-0.1800
		(0.0014)	(0.0015)	(0.1824)	(0.2011)
	≤400	0.7349	0.0780	1.2086	-0.4442
		(0.0012)	(0.0013)	(0.1750)	(0.1908)

NOTE: The sample includes white women who are living with all of their children, whose oldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth (619,499 observations)

Table A. Specification Checks and Sample Construction

Panel 1	Dependent variable: Child is a girl
Child lives with father and not with mother	-0.0501
	(0.00)
Constant	0.4888
	(0.00)
Observations	2,810,256
Panel 2	Dependent variable: Oldest child is a girl
Mother lives with all her children	0.0000
	(0.0013)
Constant	0.4871
	(0.00)
Observations	790,746
Panel 3	Dependent variable: Ever divorced
Age	0.0264
	(0.0003)
Age squared ÷ 100	-0.0305
	(0.0005)
Age at first birth	-0.0025
	(0.0001)
Dummy for high school dropout	0.0947
	(0.0012)
Observations	1,092,433

NOTES: The sample in panel 1 includes white children under age 19 (the unit of observation is a child)
The sample in panel 2 includes women meeting the set of restrictions imposed throughout the paper, except that we do not restrict to mothers living with all their children. That is, it includes white mothers who had their first birth after marriage after age 18 and before age 45, had a single first birth, and whose oldest child is under 17.
The sample in panel 3 includes white women who had their first birth after marriage, after age 18, and before age 45, and had a single first birth. Controls also include birthplace dummies (147 categories)
Robust standard errors in brackets

Figure 1 - Difference in probability of being currently married

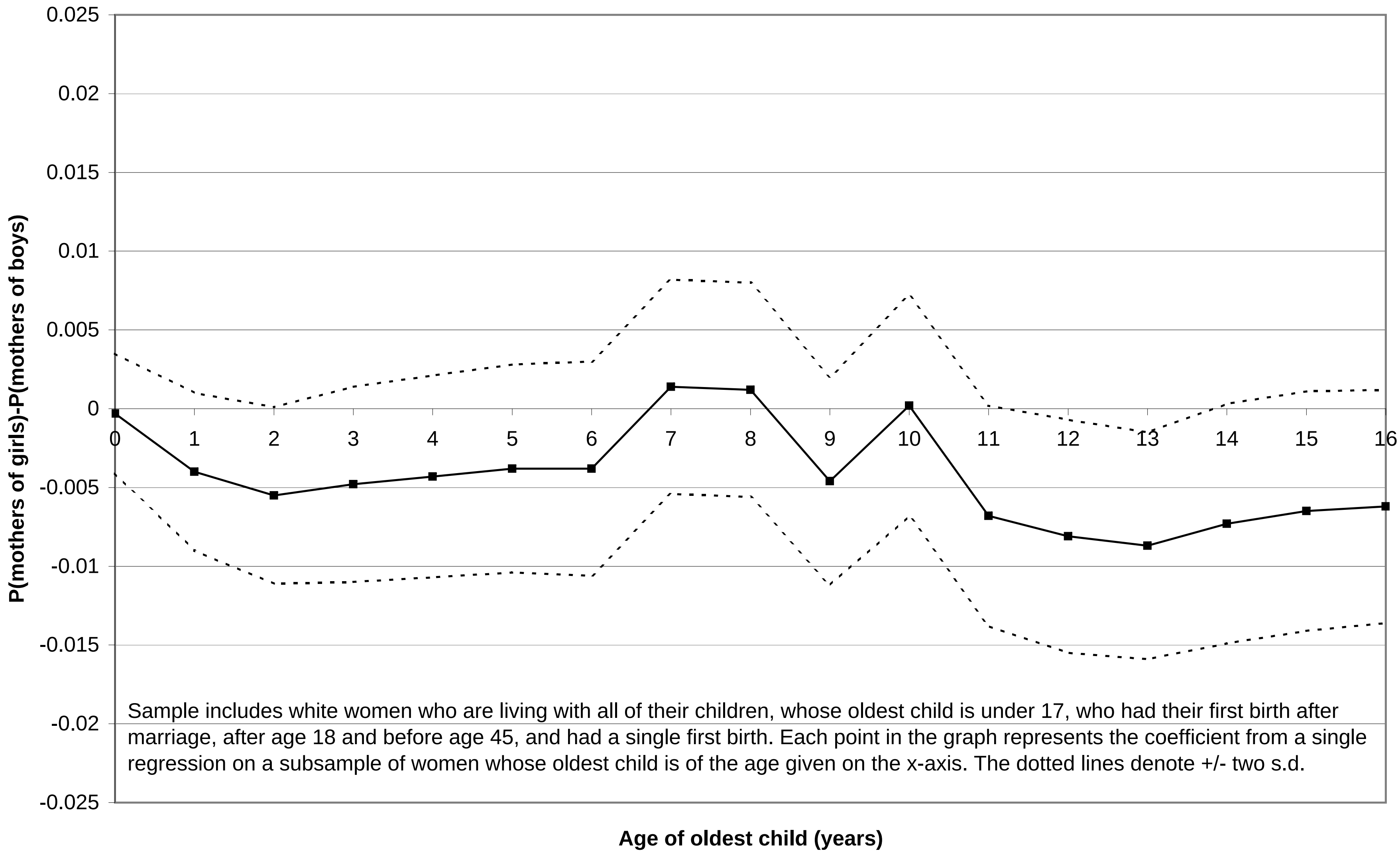

Figure 2 - Difference in probability that the first marriage is broken