Sex Preferences, Marital Dissolution and the Economic Status of Women

Kelly Bedard
Olivier Deschênes*

Department of Economics
University of California, Santa Barbara

June 2003

Abstract

The rise in the divorce rate over the past 40 years is one of the fundamental changes in American society. A seemingly ever-increasing number of women and children spend some fraction of their life in single female-headed households, leading many to be concerned about the economic circumstances of these women their and children. Estimating the cause-to-effect relationship between marital dissolution and female economic status is complicated because the same factors that increase marital instability may also affect the economic status and labor market behavior of women. We propose an instrumental variables solution to this problem based on the sex of the firstborn child. This strategy exploits the fact that the sex of the firstborn child is random and the fact that marriages are less likely to survive following the birth of girls as opposed to boys. Our IV estimates cast doubt on the contention that marital instability causes large declines in woman's economic status. Once the negative selection into divorce is accounted for, we find that women who have experienced marital dissolution have considerably higher levels of personal income and annual wages than women who remain married. At the same time we find little evidence of differential poverty rates and equivalized household incomes among ever-divorced women and never-divorced women. We further show that the higher wages of ever-divorced women mostly reflect increased labor supply intensity (hours and weeks of work) of woman who experienced marital dissolution.

* We thank Joshua Angrist, David Card, Peter Kuhn, Doug Steigerwald and Cathy Weinberger for their suggestions.

1. Introduction

According to the Center for Disease Control (July, 2002), 33% of first marriages now end in separation or divorce within 10 years. The rising incidence marital dissolution[1] has received substantial attention among social scientists and policymakers. A large body of research has documented the decline in the economic status for women who experience marital instability (i.e. Hoffman and Duncan 1988, Burkhauser et al. 1991, Smock, and Manning and Gupta 1999). Moreover, the reduction in economic status associated with divorce tends to be long-lasting, unless women remarry.

Based on these observations some analysts have concluded that marriage is a central determinant of economic status for women and their children, thereby making the case for stronger divorce laws. However, it is unclear that the negative association between marital dissolution and the economic well-being of women represents a cause-to-effect relationship: The same factors that increase the probability of divorce may also be detrimental to economic and labor market status. Therefore, a credible assessment of the impact of divorce on the economic status of women requires an exogenous determinant of marital instability.

Several concerns motivate the importance of identifying the causal relationship between divorce and the economic status of women. First, this information will contribute to a better understanding of the underlying causes of gender inequality in the United States.[2] The higher rates of poverty and economic deprivation among single female-headed families further heighten the significance of this question. Second, this has tremendous importance for the analysis of welfare and income-maintenance programs, especially in the current context of welfare reform. Welfare program recipients are disproportionately never-married or divorced mothers. Since program eligibility depends on income and labor market outcomes such as participation and/or labor supply intensity, a better understanding of the relationship between marital disruption and economic status could help improve the efficacy of these programs.

---

[1] Throughout the paper we use marital instability/disruption/dissolution and divorce interchangeably to identify women who are currently separated or have been divorced at least once.
[2] See Fuchs (1988) and Blau (1998) for overviews.

Finally, empirical studies may inform and test the predictions of theoretical models of marital formation and dissolution (Becker, Landes and Michael 1977, Lundberg and Pollak 1993).

In this paper we propose an instrumental variables (IV) approach to identify the causal effect of marital disruption on the economic status and labor market outcomes of women. The sex of the first child born during a woman's first marriage is used as an instrument for marital disruption. This exploits the fact that marriages are less likely to survive the birth of daughters than sons (Morgan, Lye and Condran 1988, Teachman and Schollaert 1989, Katzev, Warner and Acock 1994).[3] Using data from the 1980 Census, we find that the rate of marital dissolution is 4% higher for women whose firstborn child is a girl. Since firstborn sex is essentially random, we can construct a credible instrument for divorce in the population of women with at least one child born during their first marriage.[4]

We implement this instrumental variables strategy using data from the 1980 U.S. Census of Population. Cross-sectional OLS comparisons of poverty rates and equivalized household incomes among ever- and never-divorced women indicate a substantial economic disadvantage for ever-divorced women, as was previously noted. However, the evidence from our IV models cast doubt on this view. Using firstborn sex as an exogenous determinant of marital dissolution, we find that divorced women have significantly higher levels of personal income than women who remain married. At the same time, our IV estimates indicate that divorce has no significant effect on equivalized household income or poverty, although the null hypothesis of small negative effects for these two outcomes cannot be rejected.

We investigate two channels through which divorced women can raise their economic position: Labor market attachment and access to other sources of income. First, our IV results indicate that ever-divorced women earn approximately $12,000 (in 2002 dollars) per year more than their never-divorced counterparts. We show that this earnings advantage is due to greater labor supply intensity (weeks and

[3] Lundberg and Rose (2002) similarly find that the birth of a son out of wedlock accelerates the transition into marriage relative to the birth of a girl.
[4] Angrist and Evans (1998) use a similar IV strategy to estimate the impact of family size on labor supply for married women. In particular, they use the sex-mix of the first two children as an instrument for subsequent fertility.

hours of work) rather than higher labor force participation. We also show that the gap in labor supply intensity between ever- and never-divorced women is larger among women with older children.

Alternatively, divorced women can improve their economic situation by obtaining more non-labor income, welfare payments and custody transfers, or cohabitating or remarrying. We investigate this possibility by comparing non-woman income (total household income not generated by the woman) and non-wage income (a woman's total personal income minus her annual wages) across ever- and never-divorced women. Again, simple cross-sectional comparisons indicate substantially lower non-woman and non-wage income for ever-divorced women, while our IV estimates do not allow us to distinguish between zero effects or small negative effects. We also find that many divorced women with young children reside with their parents. In such cases parental income compensates for the loss of the husband's income after marital dissolution.

Overall, the IV results suggest that the marked reduction in economic status of women following marital dissolution is mostly due to the negative selection of women into divorce. Once selection is accounted for, we find little evidence of increased poverty or lower income levels among ever-divorced women. We attribute this finding to the stronger labor market attachment and higher wages of women who experience marital instability. Finally, we argue that divorce affects the labor supply of women through the intensive rather than the extensive margin.

The remainder of the paper is as follows. Section 2 describes the data. Section 3 documents the relationship between the firstborn sex and marital dissolution. Section 4 characterizes the population of ever-divorced women and demonstrates the random assignment of firstborn sex. Section 5 analyzes the relationship between marital dissolution and economic status of women. Section 6 presents a sensitivity analysis. The last section concludes.

2. Data

The data for this study are drawn from the 1980 U.S. Census Public-Use Micro Samples.[5] The 1980 Census is well suited for an analysis of marital instability and the economic well-being of women because it contains information on marital history (number of marriages and age at first marriage) and indicators of economic status (poverty, income and labor market outcomes) for a large and representative sample of women. In fact, the 1980 Census is the only large and nationally representative data set allowing this kind of analysis. The 1990 Census does not contain marital history information (only current marital status is reported), and while earlier Censuses provide marital and fertility history information, the rate of marital dissolution was too low in the 1960s and 1970s to motivate an empirical investigation.

Throughout the analysis we use a sample of white women aged 21-40 who married at least once. We focus on white women because the high rate of out-of-wedlock births among black women renders a highly selected sample.[6] Table 1 reports summary statistics and illustrates the construction of the sample used in the analysis. The first column displays summary statistics for all white women aged 21-40, irrespective of marital and fertility history. Columns 2 and 3 restrict the sample to ever-married women, and ever-married women with at least one child, respectively.[7] The entries in column 3 indicate that ever-married woman with at least one child have 2.2 children on average. By 1980, 26% of these women's first marriages had ended.

Because the Census does not identify children who have moved out of the household, we further limit the sample to women whose children all reside in her household. This restriction allows us to ascertain the sex of the oldest child.[8] Children are matched to mothers within households using the household relationship information provided in the Census. The summary statistics for this subsample of

---

[5] The data were accessed from the Integrated Public Use Microdata Series (Version 2.0). Observations with allocated age, number of marriages, current marital status, age at first marriage, number of children ever-born, relationship to the household head and sex were excluded. Families are also excluded if the oldest child has allocated values for age, sex, relationship to the household head or month of birth. Widows are also excluded. None of the results are significantly altered by these exclusions.

[6] See for example, Akerlof, Yellen and Katz (1996) who report that in 1980-84, 57% of black births were out-of-wedlock compared to only 12% of white births.

[7] Throughout we consider women whose first marriage began when they were 17-26 years of age. This includes 90% of all first marriages.

[8] To avoid confounding firstborn sex and family size we also exclude women with firstborn twins. However, the occurrence of firstborn twins is so small that their inclusion or exclusion is immaterial.

558,808 mothers are reported in column 4 of Table 1. For this subsample, 21% of first marriages ended in divorce. The lower incidence of marital instability among this group reflects the fact that they are less likely to have had their first child before their first marriage began, which makes their first marriage more likely to survive (Bronars and Grogger 1994).

Column 5 further restricts the sample to women whose oldest child is under the age of eighteen, and whose children all live in her household. This requirement is important because youth are progressively more likely to move out of their parent's home as they age. Finally, the Census does not distinguish between biological, adopted and stepchildren. In an attempt to isolate biological children born during the first marriage, we limit the sample to women whose first child is born within the first five years of her first marriage. Combined, these restrictions allow us to identify the sex of the firstborn child within the first marriage. The summary statistics for this final sample of 465,595 women are reported in column 6. Except where noted otherwise, the empirical analysis in the remainder of the paper uses the sample described in column 6. For simplicity, we refer to this group as 'ever-married mothers'. The incidence of marital dissolution for this group is 20%. The similarity of the averages in columns (1)-(6) indicate that the sample of ever-married mothers is mostly representative of the population of white women aged 21-40. Nevertheless, section 6 examines the sensitivity of our results to these restrictions.

We use five indicators of economic status: equivalized household income (total household income divided by the poverty line[9]), poverty (=1 if the household's total income places them below the poverty line), non-woman income (total household income minus the total income of the woman), personal income (the woman's total income) and the woman's annual wage. Since the family unit is not always a well-defined concept following marital dissolution, due to cohabitation, we focus on household-based measures of income instead of the family-based measures reported in the Census.[10]

The average economic status of the women in our sample is reported in the bottom panel of Table 1. The average level of household income among ever-married mothers is 312.5 times the poverty line.

[9] We translate the family poverty cut-offs, based on the number of adults and children, to households.
[10] We obtain very similar results when we use the family-based measures of income reported in the Census.

The poverty rate is 7% and average personal income is $9,437, almost all of which is labor market earnings ($8,245). Thus, non-labor sources of income such as welfare payments, child support and alimony do not appear to be important income sources for the population of ever-married mothers. On average, the largest income component is non-woman income at $40,185. Combined, these income levels imply an average household income of approximately $50,000.

3. The relationship between firstborn sex and marital dissolution

The fact that the sex composition of offspring, in families with at least 2 children, affects subsequent fertility is often interpreted as evidence of parental preference for mixed sex composition among their offspring (i.e. Ben-Porath and Welch 1976, Leung 1991, Angrist and Evans 1998). Others have suggested that couples may have a preference for firstborn boys (Williamson 1976). However, parental sex preferences may also influence marital stability. For example, some researchers have argued that marriages are less likely to survive the birth of daughters than sons (Morgan, Lye and Condran 1988, Teachman and Schollaert 1989, Katzev, Warner and Acock 1994). Assuming that firstborn sex is random (evidence for this is provided below) and that there is a systematic relationship between divorce and the sex of the firstborn child, firstborn sex may provide an exogenous source of variation in the probability that a first marriage ends among families with at least one child.

Table 2 investigates the relationship between firstborn sex and marital instability. We present several reduced-form estimates that confirm that firstborn girls increase the probability that the marriage ends in divorce. We also explore the potential sources of heterogeneity in the relationship between marital instability and firstborn sex. Columns 1 and 2 report the unadjusted differences and associated F-statistics testing the null hypothesis that firstborn sex has no effect on marital instability and columns 3 and 4 report the regression-adjusted differences and associated F-statistics.[11] Finally, column 5 displays the number of observation used in each model.

Panel A reports the overall effect of firstborn sex on the probability of marital disruption. The point estimate is 0.008 (std error=0.001), indicating that in the population of ever-married mothers, a firstborn born girl increases the probability that the first marriage ends by 0.8 percentage points. This translates into a 4% higher divorce rate for women with firstborn girls relative to firstborn boys, given an average divorce rate of 20% for all first marriages. Adjusting for observable characteristics does not alter the estimated effect of a firstborn girl on marital disruption, as should be expected if firstborn sex is randomly assigned. The large F-statistics (50.2 and 46.1 for the unadjusted and adjusted models, respectively) confirm the importance of firstborn child's sex on marital instability.

Panels B, C and D investigate the potential sources of heterogeneity in the relationship between marital instability and firstborn sex. In panel B we estimate the models separately by education level (dropout, high school graduate, some college and college graduate). While the effects are more pronounced for high school dropouts, firstborn sex significantly affects the probability of divorce for all education groups except college graduates (the adjusted F-statistics are respectively 18.1, 13.7, 15.9 and 3.4). Thus, the results reported in panel A are not driven by a single group of women, at least in terms of educational attainment.

Panel C allows for differential effects by mother's age at first marriage. In particular, we estimate the models separately if the mother married before or after age of 20. While the effect of first offspring sex on marital disruption is stronger for mothers who married at younger ages, the F-statistics indicate a significant relationship for both age groups. Panel D similarly allows for differential effects by age at first birth. We break the sample into women whose first birth occurred before or after her twenty-second birthday. Again, the contrasts are significant for both groups, but the effect of a firstborn girl is stronger among women who were younger when they had their first child.

[11] Unless noted otherwise all the models in this paper control for the following characteristics: quadratics in age, age at first marriage, age at first birth and education, unrestricted state of birth and residence dummies, a dummy for SMSA status and interactions between education and the other continuous explanatory variables.

Finally, we also note that this relationship is not specific to the 1980 Census. We also examined the effect firstborn sex on the incidence of marital instability using data from the 1960 and 1970 Censuses. The point estimates for 1960 and 1970 are 0.006 and 0.004, respectively, with standard errors of 0.002 and 0.001. These effects translate into 3-7% higher divorce rates for women with firstborn girls, relative to the average divorce rates for these years. An appendix containing these results is available from the authors upon request.

4. The random assignment of firstborn sex and characteristics of ever-divorced women

In the neoclassical theory of marriage (Becker 1973, 1974), marital gains are derived from specialization within the household and hence depend on the woman's potential earnings capacity relative to her husbands, which is determined by the marital matching process (i.e. Burdett and Coles 1997). Therefore the characteristics of husbands and wives whose marriages end will differ from the average characteristics of husbands and wives whose marriages continue. As a result, cross-sectional comparisons of labor market outcomes across never- and ever-divorced women may be confounded by omitted variables bias.

The left panel Table 3 provides some evidence suggesting that marital disruption is affecting a nonrandom subset of the ever-married mother population. Columns 1 and 2 report the average marital history, fertility and socioeconomic characteristics of never- and ever-divorced women, respectively. Column 3 reports the mean differences and their associated standard errors. These entries provide clear evidence that marital instability is not randomly assigned. In particular ever-divorced women were younger when they married for the first time, were younger when their first child was born and are less educated. All differences reported in column 3 are statistically significant at the 5% level. These substantial differences highlight the need for a credible research design capable of addressing the differential selection into never- and ever-divorced groups.

In contrast to the important differences in the characteristics of women across divorce status, there is no evidence of systematic differences in the observable characteristics across women with firstborn girls and boys. This is evidenced in the right panel of Table 3. Columns 4 and 5 report the average marital history, fertility and socioeconomic characteristics of ever-married mothers with firstborn girls and boys, respectively, and column 6 reports the differences in means and their associated standard errors. None of the differences reported in column 6 are statistically significant at the conventional level. Moreover, the differences are all very small relative to the sample averages reported in the sixth column of Table 1. This is exactly what should be expected if firstborn sex is randomly assigned in the population of ever-married mothers. Of particular importance, age at first marriage, age at first birth and education, three strong predictors of marital instability (as evidenced in column 3) are balanced on the basis of the firstborn sex. Overall, the results reported in the right panel of Table 3 support the claim that firstborn sex is randomly assigned within the population of ever-married mothers, at least in terms of observable characteristics. In the next sections, we use an instrumental variables strategy based on this fact to analyze the impact of divorce on the economic status of women.

5. The effect of marital dissolution on women economic status

A. OLS estimates

We begin the empirical analysis by using OLS regressions to estimate the relationship between women economic status and marital instability.[12] Let Y_i denote a measure of economic status for woman i:

$$Y_i = \alpha + \beta D_i + X_i \gamma + \varepsilon_i \quad (1)$$

As previously stated, we focus on five indicators of economic status: equivalized household income, poverty, non-woman income, personal income and wages. In addition, we also analyze the labor supply determinants of annual wages: employment, weeks worked last year and hours worked per week. The variable D_i is a dummy variable indicating that the woman's first marriage dissolved, X_i denotes observable characteristics, and ε_i represents the unobservable determinants of economic status. The parameter of interest is β; the causal effect of marital instability on economic status. In all models, X_i includes quadratics in age, age at first marriage, age at first birth and years of education, a dummy for

[12] For the binary outcomes, poverty and employment, the OLS models correspond to linear probability models. In all cases the probit estimates of the marginal effects are nearly identical.

SMSA status, unrestricted state of birth and state of residence dummies, as well as interactions between years of education and all the other continuous explanatory variables.

The first column in Table 4 reports the OLS estimates of β in equation (1). These estimates show the negative cross-sectional association between economic status and marital dissolution, as others have documented. All the effects in column 1 are very precisely estimated. Equivalized household income is 53 income equivalent points (17%) lower and the poverty rate is 12 percentage points higher for ever-divorced mothers. This is entirely explained by the large reduction in non-woman income following divorce. As column 1 shows, on average ever-divorced women have higher personal income than never-divorced mothers, a fact mostly attributable to their higher labor market earnings. As indicated by the last rows in column 1, the higher wages arise from the stronger labor market attachment of ever-divorced women. Overall, these estimates are qualitatively similar to those in other studies (i.e. Johnson and Skinner 1986, Duncan and Hoffman 1985 and 1989, Burkhauser et al. 1991, Holden and Smock 1991, Smock 1993, Smock et. al. 1999).

B. TSLS estimates

While there is well-documented evidence of significant marital status effects in models of female economic status and labor supply, like in column 1 of Table 4, the causal interpretation of these estimates in various contexts has been questioned in recent years (Korenman and Neumark 1992, Smock, Manning and Gupta 1999, Krashinsky 2002). In particular, the causal interpretation rests on the assumption that unobservables do not confound the marital status effect. This seems very unlikely. Table 3 clearly documents substantial differences in observable characteristics across never- and ever-divorced women. This suggests that the populations of never- and ever-divorced women may also differ along unobservable dimensions, in particular, unobserved determinants of economic status and labor supply (Altonji, Elder and Taber 2000). In other words, D_i and ε_i may be correlated in equation (1). In this case, the OLS estimates of β reported in Table 4 will be confounded by omitted variable bias.

We propose an instrumental variables solution to this problem using the sex of the firstborn child as an exogenous determinant of divorce. This identification strategy rests of the assumption that firstborn sex influences indicators of economic status like poverty and income only through its effect on marital stability, so that firstborn sex can be rightfully excluded from models like (1). Since firstborn sex is essentially random, as evidenced in Table 3, this assumption appears reasonable. Moreover, as reported in Table 2, firstborn sex is an important determinant of marital disruption with an F-statistic of 46.1. This approach may therefore be helpful in identifying the causal effect of divorce on economic status.

To exploit the randomness embodied in the sex of the firstborn child, we begin by using a single indicator for firstborn girls as an instrument for marital dissolution. More specifically, we estimate the parameters of equation (1) using TSLS based on the following the first-stage equation for divorce:

$$D_i = \pi_0 + \pi_1 G_i + X_i \pi_2 + v_i. \tag{2}$$

Estimates of π_1 have already been reported in Table 2.[13] Column 2 in Table 4 reports the TSLS (or "Wald") estimates of the effect of marital instability on the determinants of economic status when the control variables, X_i, are excluded from the first-stage and outcome equations. Column 3 in Table 4 similarly reports the TSLS estimates when the control variables are included in the models. While the model in (1) is written as a homogeneous treatment effect model, the TSLS estimates in Table 4 can be interpreted as the LATE specific to the instrument firstborn sex (Imbens and Angrist 1994, Angrist and Imbens 1995). Under this interpretation, the TSLS estimate of β is the average treatment effect in the population of mothers whose marital status is changed by the sex of their firstborn child.

The entries in columns 2 and 3 point to two main findings: First, there is no evidence of large systematic differences in the economic status of ever- and never-divorced women once the negative selection of women into divorce is accounted for. Secondly, the higher level of personal income and wages for ever-divorced women is attributable to greater labor supply intensity rather than labor market participation. These conclusions are supported by two sets of results.

[13] All reduced-form estimates are reported in an appendix and are available from the authors upon request.

First, the TSLS estimates indicate that personal income and annual wages are $13,042 and $11,378 higher on average for ever-divorced women (both differences are statistically significant at the 5% level). This contributes greatly to improving the economic position of divorced women. The evidence on poverty, equivalized income and non-woman income is more mixed. In all but one model we fail to reject the null hypothesis of zero effects, but since the TSLS estimates for these outcomes are imprecise, we cannot reject the null hypothesis of small negative effects. For example, the estimated difference in poverty rates is 0.087, with a standard error of 0.093. Despite the large standard errors, one surprising results in column 3 is the positive effect of divorce on equivalized income and non-woman wage. We investigate some explanations for this result in Table 5.

Second, the higher wages of ever-divorced women are mostly attributable to greater labor supply intensity. The TSLS estimates in column 3 reveal that the average ever-divorced woman works 24 more weeks per year than the average never-divorced women, and that this difference is statistically significant at the conventional level. The corresponding OLS estimate is 2.5 times smaller. However, the TSLS estimates for the probability of employment and hours worked per week are insignificant at the conventional level. In both cases, the IV estimates cannot distinguish between the null of zero effect or the OLS estimates reported in column 1 (those are 0.188 (employment rate) and 9.663 (hours per week)). Taken as whole, the evidence in columns 1-3 suggest that the OLS estimates overstate the effect of marital instability on the incidence of poverty and understate its effect on personal income, wages, non-woman income and household equivalent income. This is consistent with the selection of more disadvantaged women into divorce.

The last column of Table 4 adds total fertility and current marital status (=1 if the first marriage ended and the respondent is currently married) to the set of control variables. Controlling for total fertility alone does not alter the TSLS estimates reported in Table 4. This finding is consistent with the evidence in Table 2, which shows that sex of the firstborn is orthogonal to total fertility, as Angrist and Evans (1998) also found. Controlling for the current marital status of ever-divorced women is more problematic since there is non-random selection into remarriage.[14] Thus the balancing property of firstborn sex documented in Table 3 is lost once we condition on current marital status of ever-divorced women. Consequently, the TSLS estimates reported in column 4 should be interpreted with caution. Adding a control for the current marital status of ever-divorced women raises the estimated effect of marital dissolution on all outcomes considered in Table 4, although not significantly. The higher point estimates are largely driven by a weaker first stage relationship between firstborn girls and marital disruption once remarriage is controlled for. The coefficient on firstborn sex falls from 0.008 to 0.004.

C. Allowing the effects to vary by firstborn age

The effect of marital dissolution on the economic welfare of mothers may depend on the age of their firstborn for a series of reasons.[15] First, independently of parental sex preferences, firstborn sex cannot have an immediate effect on the probability of divorce. Second, strong labor market attachment is more costly and difficult among divorced mothers with younger children (relative those with older children). Finally, since husband's income grows over time (because of the increasing age-earnings profile in our sample) the decline in non-woman income following divorce will increase with the age of firstborn, since parents with older children tend to be older themselves. At the same time, divorced women with younger children may be more likely to relocate with their own parents following divorce, which could alleviate the loss of their husband's income.[16]

We explore these possibilities in Table 5 by estimating the models on samples of women whose firstborn is younger or older than 12 years old. For comparison, column 1 reports the unconditional results from column 3 in Table 4. The OLS estimates reported in the top panel of Table 5 confirm the finding of a negative cross-sectional association between marital disruption and economic status for

[14] For example, ever-divorced mothers who remarry are older and are better educated than those who do not.
[15] We elected to condition on firstborn age because women with firstborn children of the same age have been at risk of marital instability (at least the part caused by firstborn sex) for the same period of time. Stratifying the analysis by woman's age would tend to confound this kind of effect. Nevertheless, there is a strong correlation between mother's age and firstborn age.
[16] Card and Lemieux (2000) show that fraction of youths living with their parents is an important determinant of the poverty trends among youths in Canada and the U.S.

women illustrated in Table 4. The point estimates in columns 2 and 3 are remarkably similar to those in column 1. Thus the OLS results appear largely independent of firstborn age.

In the bottom panel of Table 5 we report the TSLS estimates following the same specification as column 3 in Table 4. For the two samples we also report the F-statistics from the first-stage relationship between marital dissolution and firstborn sex. In both cases the F-statistics are large (18.9 and 29.7, respectively), with the strongest effect for the group of women whose firstborn child is aged 12 or older. While sometimes imprecise, the TSLS estimates in columns 2 and 3 indicate that economic consequences of marital instability depend to an important extent on the age of the oldest child.

First, note that the positive TSLS effect of divorce on personal income and wages reported in Table 4 is mostly concentrated in the group of women whose oldest child is at least 12. The personal income and wage differentials are \$22,931 and \$19,387, respectively, and are precisely estimated. Among women with younger firstborn children the TSLS estimates for personal income and wages are small and not statistically significant. This reflects the fact that the labor supply intensity of divorced mothers increases as their children age, relative to never-divorced mothers: The estimated effects on weeks worked and hours worked per week are 38.4 and 22.6, respectively (standard errors=11.1 and 8.3).

Next we turn to non-woman income. Again, the TSLS estimates differ greatly in the two samples defined by the age of the oldest child. The estimated TSLS differences across divorce status are \$26,982 for the group with young children and -\$19,491 for the group with older children. Both point estimates are relatively imprecise, and must be interpreted accordingly. Nevertheless, the sign reversal can be accounted for by the entries reported in the last two rows of Table 5. There we report the fraction of ever-divorced mothers who reside with their parents, and the proportion who remarried. The negative relationship in the TSLS estimates between non-woman income and age of the oldest child can be explained by two phenomena. First, divorced mothers with young child are more likely to reside with their parents (by 5.6 percentage points), as shown in the last row of Table 5. Since this group of women is relatively young, it is plausible that parents of divorced mothers have higher earnings than the husbands of mothers who have remain married, which would tend to increase non-woman income among divorced mothers. Second, among mothers with older children the negative TSLS estimate on non-woman income can be attributed to the increasing age-earnings profile of husband of never-divorced women, which makes the cost of divorce higher in terms of non-woman income. Also, the proportion of divorced mothers living with their parents is smaller for this group.

Together, the fact that ever-divorced women with older children have higher personal income and wages but lower non-woman income explains why the cost of marital dissolution—as measured by higher poverty rates and lower equivalized household income—is larger for women with older children. Again, this interpretation is limited by the fact that the effects on poverty, equivalized income and non-woman income are imprecisely estimated. However, the positive and significant differences in personal income, annual wages and labor supply intensity contribute to improving the economic status of divorced women.

6. Sample selection issues

So far the analysis has been based on a sample of ever-married mothers whose children all reside in her household. This restriction is dictated by the fact the Census does not identify non-residential children. Children may move out of their mother's household following the breakup of a family, or for other reasons (mortality, college attendance, foster homes, etc). This introduces a missing-data problem: the characteristics of children who reside outside of the mother's household are not observed. This missing data problem is troublesome if the probability that a child does not reside in their mother's household is partly determined by their sex. For example, if boys are less likely than girls to live with their mother after a divorce has occurred our results may be biased in favor of finding that women with firstborn girls are more likely to experience divorce.

We address this issue in two ways. First, we examined the living arrangements of children from disrupted families using data from the 1980, 1985 and 1990 June CPS.[17] The June CPS contains complete marital and fertility history information. In particular, information on the living arrangements of all children of a mother is available. Of all firstborn children under the age of eighteen from divorced

[17] For comparability the sample definitions are identical to the restrictions placed on the Census sample.

families: 87% live with their mother, 8% live with their father and 5% live with neither their mother nor father. Most importantly for our purposes, there is little difference in the sex composition of firstborn children from divorced families living outside of their mother's household: Conditional on residing outside of the mother's household, 48% of firstborn children are girls compared to 49% in the population at large (i.e. independently of parental marital status or living arrangements). Firstborn girls are therefore over-represented by approximately 1-percentage point in our sample of ever-divorced mothers.

While it is unclear whether such a small difference could systematically bias our estimates, we nonetheless explore the impact of reintroducing the excluded part of the sample (column 3 in Table 1) by randomly assigning the sex of the firstborn child for the households where that variable cannot be defined because of the structure of the Census. This reintroduces in the sample women whose first birth was before their first marriage, and whose first birth was outside the first 5 years of the first marriage. It also includes women whose oldest child is over the age of 17 and whose children do not all live in her household. This reintroduces and additional 196,626 women to the sample.

For these 196,626 women we randomly impute undefined firstborn sex in a variety of ways. First, we simply assign 49% of women with missing firstborn child sex information firstborn girls and 51% of these women firstborn boys; these proportions match the sample averages in the Census. To check the sensitivity of the results to this choice, we also do similar imputations with 48, 47 and 46% firstborn girls. The results of the sensitivity analysis yields TSLS estimates similar to those reported in Tables 4 and 5. Thus the conclusions of Tables 4 and 5 are not specific to our baseline sample.[18]

7. Conclusion

The connection between marital instability and the economic well-being of women and children is a topic of great importance to social scientists and policymakers alike. Despite the significance of this question for many practical and theoretical debates, there is still considerable uncertainty regarding the cause-to-effect relationship between marital instability and the economic status of women. This uncertainty

[18] These results are reported in an appendix that is available from the authors upon request.

reflects the difficult task of identifying the causal relationship: The same factors that contribute to increasing the probability of divorce may also be detrimental to the economic well-being of women.

In this paper we present evidence based on an instrumental variables strategy. Our IV estimates are derived using firstborn sex as an exogenous determinant of marital instability in the population of ever-married women with at least one child. This instrument exploits the fact that marriages are less likely to survive following the birth of girls as opposed to boys. In the sample we consider, families with firstborn girls are 4% more likely to experience marital dissolution than families with firstborn boys. Since firstborn sex is essentially random, as evidenced in this paper, this approach may be helpful in untangling the "true" economic consequences of marital dissolution from the confounding effect of non-random selection of women into divorce.

Our OLS estimates of the economic consequences of marital dissolution are consistent with the previous literature in that they indicate a substantial economic disadvantage for divorced women. In particular, divorce is associated with higher poverty rates, lower equivalized household income and lower non-woman income. While similar findings have been interpreted by some analysts as indicating that marital status is an important causal determinant of female economic well-being, our IV estimates cast serious doubt on the interpretation. Once the negative selection of women into divorce is accounted for, we find that ever-divorced mothers have substantially higher levels of personal income and annual wages than never-divorced mothers, largely because they supply labor more intensely (more hours per week and weeks per year). At the same time, our IV estimates reveal little difference in the level of non-woman income among ever- and never-divorced mothers. This result is partly explained by the fact that some divorced mothers with young children relocate with their parents, thereby raising the amount of non-woman income in their household.

Since a woman's total income is the sum of her personal income and non-woman income, the estimated effect of marital disruption on equivalized household income and poverty are proportional to the estimates effects for personal and non-woman income. Consequently, our IV evidence for poverty and equivalized income is not consistent with the contention that marital dissolution causes large declines in economic status. In all models we fail to reject the null hypothesis of no effect of marital instability on poverty and equivalized income. However, since the IV estimates on these outcomes are sometimes imprecise, we cannot always distinguish between the hypotheses and small negative and zero effects. That being said, taken as a group our results clearly suggest that the most of the economic costs associated with marital dissolution can be attributed to the negative selection of women into divorce.

References

Akerlof, George A., Janet L. Yellen and Michael L. Katz. 1996. "An Analysis of Out-of-Wedlock Childbearing in the United States" *Quarterly Journal of Economics* 111(2): 277-317.

Altonji, Joseph G., Todd E. Elder and Christopher R. Taber. 2000. "Selection on Observed Variables: Assessing the Effectiveness of Catholic Schools," NBER Working Paper 7831.

Angrist, Joshua D. and Guido W. Imbens. 1995 "Two-Stage Least Squares Estimation of Average Causal Effects in Models with Variable Treatment Intensity," *Journal of The American Statistical Association*, 90: 431-442.

Angrist, Joshua D. and William N. Evans. 1998. "Children and Their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size," *American Economic Review* 88(3): 450-77.

Becker, Gary S. 1973. "A Theory of Marriage: Part I," *Journal of Political Economy* 81(4): 813-46.

Becker, Gary S. 1974. "A Theory of Marriage: Part II," *Journal of Political Economy* 82(2): S11-26.

Becker, Gary S., Elisabeth M. Landes and Robert T. Michael. 1977. "An Economic Analysis of Marital Instability," *Journal of Political Economy* 85(6): 1141-88.

Ben-Porath, Yoram and Finis Welch. 1976. "Do Sex Preferences Really Matter?" *Quarterly Journal of Economics* 90(2): 285-307.

Blau, Francine D. 1998. "Trends in the Economic Well-Being of Women 1970-1995" *Journal of Economic Literature* 36(1): 112-165.

Bronars, Stephen G. and Jeff Grogger. 1994. "The Economic Consequences of Unwed Motherhood: Using Twin Births as a Natural Experiment," *American Economic Review* 84(5): 1141-56.

Burkhauser, Richard, Greg J. Duncan, Richard Hauser, and Roland Bersten. 1991. "Wife or Frau, Women do Worse: A Comparison of Men and Women in the United States and Germany After Marital Dissolution," *Demography* 28(3): 353-60.

Burdett, Ken and Melvin G. Coles. 1997. "Marriage and Class," *Quarterly Journal of Economics* 112(1): 141-68.

Card, David and Thomas Lemieux. 2000. "Adapting to Circumstances: The Evolution of Work, School, and Living Arrangements among North American Youth," in David G. Blanchflower and Richard Freeman (eds.) Youth Employment and Joblessness in Advanced Countries, University of Chicago Press.

Center for Disease Control 2002. "Cohabitation, Marriage, Divorce, and Remarriage in the United States," Available online at: http://www.cdc.gov/nchs/releases/02news/div_mar_cohab.htm

Duncan Greg J. and Saul D. Hoffman. 1985. "A Reconsideration of the Economic Consequences of Marital Dissolution," *Demography* 22(4): 485-97.

Fuchs, Victor R. 1988. *Women's Quest for Economic Equality*, Harvard University Press, Cambridge Massachusetts.

Hoffman, Saul D., and Greg J. Duncan. 1988. "What are the Economic Costs of Divorce," *Demography* 25(4): 641-45.

Holden, Karen C. and Pamela J Smock. 1991. "The Economic Costs of Marital Dissolution: Why Do Women Bear a Disproportionate Cost?" *Annual Review of Sociology* 17: 51-78.

Imbens, Guido W. and Angrist, Joshua D. 1994 "Identification and Estimation of Local Average Treatment Effects," *Econometrica*, 62: 467-475.

Johnson. William R. and Jonathan Skinner. 1986. "Labor Supply and Marital Separation," *American Economic Review* 76(3): 455-69.

Katzev, Aphra R., Rebecca L. Warner and Alan C. Acock. 1994. "Girls or Boys? Relationship of Child Gender to Marital Instability," *Journal of Marriage and the Family* 56(2): 89-100.

Korenman, Sanders and David Neumark. 1992. "Marriage, Motherhood, and Wages," *Journal of Human Resources* 27(2): 233-55.

Krashinsky, Harry A. 2002 "Do Marital Status and Computer Usage Really Change the Wage Structure?" mimeo, University of Toronto.

Leung, Siu Fai. 1991. "A Stochastic Dynamic Analysis of Parental Sex Preferences and Fertility," Q*uarterly Journal of Economics* 106(4): 1063-88.

Lundberg, Shelly and Robert A. Pollak. 1993. "Separate Spheres Bargaining and the Marriage Market," *Journal of Political Economy* 101(6): 988-1010.

Lundberg, Shelly and Elaina Rose. 2002. "The Effects of Sons and Daughters on Men's Labor Supply and Wages," *Review of Economics and Statistics* 84(2): 251-68.

Morgan, Philip S., Diane N. Lye and Gretchen A. Condran. 1988. "Sons, Daughters, and the Risk of Marital Disruption," *American Journal of Sociology* 94(1): 110-29.

Smock Pamela J. 1993. "The Economic Costs of Marital Disruption for Young Women over the Past Two Decades," *Demography* 30(3): 353-71.

Smock, Pamela J., Wendy D. Manning, and Sanjiv Gupta. 1999. "The Effect of Marriage and Divorce on Women's Economic Well-Being," *American Sociological Review* 64(4): 794-812.

Teachman, Jay D. and Paul T. Schollaert. 1989. "Gender of Children and Birth Timing," *Demography* 26(3): 411-22.

Williamson, Nancy E. 1976. *Sons or Daughters: A Cross-Cultural Survey of Parental Preferences*, Beverly Hills, CA: Sage.

Table 1: Descriptive Statistics and Construction of the Samples from the 1980 Census

	(1) All Women	(2) Ever-Married	(3) Ever-Married with Children	(4) All Children Live in Household	(5) Oldest Child Aged 17 or Younger	(6) 1st Child Born within 5 Years of 1st Marriage
Marital History						
First Marriage Ended	--	0.26	0.26	0.21	0.21	0.20
		(0.44)	(0.44)	(0.41)	(0.41)	(0.40)
Age at First Marriage	--	20.24	19.93	20.04	20.10	20.03
		(2.24)	(2.15)	(2.14)	(2.14)	(2.13)
Fertility						
Firstborn Girl	--	--	--	0.49	0.49	0.49
				(0.50)	(0.50)	(0.50)
Children Ever-Born	1.42	1.69	2.18	2.07	2.03	2.08
	(1.39)	(1.32)	(1.08)	(0.96)	(0.93)	(0.94)
Age at First Birth	--	--	--	22.50	22.62	22.17
				(3.21)	(3.21)	(2.68)
Socioeconomic Characteristics						
Age	29.60	30.38	31.38	30.94	30.61	30.53
	(5.58)	(5.35)	(5.15)	(5.00)	(4.84)	(4.89)
Years of Education	12.97	12.90	12.66	12.76	12.80	12.74
	(2.41)	(2.19)	(2.10)	(2.09)	(2.09)	(2.02)
Urban	0.68	0.66	0.64	0.65	0.65	0.64
	(0.47)	(0.47)	(0.48)	(0.48)	(0.48)	(0.48)
Economic Status						
Equivalized Household	362.5	357.4	321.2	320.2	318.9	312.5
Income	(220.7)	(212.6)	(186.5)	(181.8)	(181.9)	(176.2)
Poverty	0.07	0.06	0.07	0.07	0.07	0.07
	(0.25)	(0.24)	(0.26)	(0.25)	(0.25)	(0.25)
Non-Woman Income	37,213.1	38,811.5	40,402.2	40,808.8	40,335.0	40,184.9
	(29,363.5)	(27,712.7)	(28,279.7)	(27,912.5)	(27,717.6)	(27,627.7)
Woman's Income	13,370.8	12,170.4	10,130.3	9,768.7	9,643.9	9,436.8
	(14,020.6)	(13,737.3)	(12,982.3)	(12,823.7)	(12,763.7)	(12,537.9)
Wages	12,116.7	10,952.1	8,828.8	8,549.0	8,433.2	8,245.0
	(13,243.0)	(12,832.8)	(11,860.2)	(11,701.0)	(11,638.5)	(11,411.3)
Sample Size	1,188,126	854,460	662,204	558,808	535,887	465,595

The baseline sample includes all women who are currently aged 21-40. Each column progressively restricts the sample as defined in the column header. Ever-married is defined as the first marriage beginning when the respondent was 17-26 years old. Standard deviations are reported in parentheses. All dollar figures are in 2002 constant dollars.

Table 2: Effect of Firstborn Sex on Marital Dissolution

Dep=First Marriage Ended	(1) Unadjusted	(2) F-Statistic	(3) Reg-Adjusted	(4) F-Statistic	(5) Observations
<u>(A) Overall Effect</u>					
Firstborn Girl	0.008	50.2	0.008	46.1	465,595
	(0.001)		(0.001)		
<u>(B) By Education Level</u>					
<12 years	0.018	20.8	0.016	18.1	51,981
12 Years	0.006	15.4	0.006	13.7	247,053
13-15 Years	0.011	18.6	0.009	15.9	102,126
16+ Years	0.004	2.7	0.004	3.4	64,435
<u>(C) By Age at First Marriage</u>					
<20 Years Old	0.011	34.4	0.010	29.7	216,822
20+ Years Old	0.006	17.2	0.006	16.6	248,773
<u>(D) By Age at First Birth</u>					
<22 Years Old	0.012	39.2	0.011	36.3	207,584
22+ Years Old	0.005	11.9	0.005	12.3	258,011

The sample is defined as in column 6 in Table 1. The models in (3) and (4) include quadratics in age, age at 1st marriage, age at 1st birth and years of education, interactions between education, age, age at 1st marriage and age at 1st birth, as well as unrestricted dummies for state of birth, state of residence and residence in a SMSA. Heteroskedasticity-robust standard errors are reported in parentheses.

Table 3: Difference in Means, by Divorce Status and Firstborn Sex (Ever-Married Mothers)

	(1)	(2)	(3)	(4)	(5)	(6)
	By Divorce Status			By Firstborn Sex		
	Never-divorced	Ever-divorced	Difference	Firstborn Girl	Firstborn Boy	Difference
Marital History						
First Marriage Ended	---	---	---	0.203	0.195	0.008
				(0.402)	(0.396)	(0.001)
Age at First Marriage	20.207	19.310	-0.897	20.025	20.031	-0.006
	(2.142)	(1.910)	(0.007)	(2.126)	(2.131)	(0.006)
Fertility						
Firstborn Girl	0.485	0.498	0.013	---	---	---
	(0.500)	(0.500)	(0.002)			
Number of Children	2.114	1.964	-0.150	2.086	2.082	0.004
	(0.935)	(0.943)	(0.003)	(0.943)	(0.933)	(0.003)
Age at First Birth	22.406	21.204	-1.202	22.163	22.171	-0.008
	(2.681)	(2.432)	(0.009)	(2.677)	(2.677)	(0.008)
Socioeconomic Characteristics						
Age	30.477	30.721	0.244	30.529	30.521	0.008
	(4.961)	(4.609)	(0.017)	(4.891)	(4.897)	(0.014)
Years of Education	12.823	12.395	-0.428	12.741	12.736	0.005
	(2.035)	(1.912)	(0.007)	(2.016)	(2.021)	(0.006)
Urban	0.633	0.679	0.047	0.643	0.641	0.002
	(0.482)	(0.467)	(0.002)	(0.479)	(0.480)	(0.001)
Sample Size	373,067	92,528	465,595	227,218	238,377	465,595

The sample is defined as in column 6 in Table 1. The standard deviations are reported in parentheses except in columns 3 and 6 the entries in parentheses are heteroskedasticity-robust standard errors.

Table 4: OLS, Wald and TSLS Estimates of Women Economic Status and Labor Supply Models

Dependent variable:	(1) OLS	(2) WALD	(3) TSLS	(4) TSLS
Equivalized Household Income	-53.1	140.3	136.0	220.0
	(0.6)	(67.9)	(67.4)	(113.5)
Poverty	0.119	0.068	0.087	0.143
	(0.001)	(0.088)	(0.093)	(0.146)
Non-Woman Income	-20,856.1	6,476.6	5,161.3	4,840.9
	(104.0)	(10,079.6)	(10,147.6)	(16,161.1)
Woman's Income	8,680.8	13,801.8	13,041.5	20,969.8
	(50.1)	(4,324.0)	(4,487.8)	(7,185.8)
Wages	5,779.6	11,995.6	11,377.5	18,983.6
	(46.1)	(4,048.9)	(4,214.8)	(6,834.0)
Working for Pay	0.188	0.192	0.148	0.279
	(0.002)	(0.169)	(0.179)	(0.281)
Weeks per Year	10.163	25.869	24.492	41.983
	(0.081)	(8.054)	(8.450)	(13.864)
Hours per Week	9.663	10.556	9.626	16.348
	(0.067)	(6.423)	(6.784)	(10.622)
Controls for total fertility and current marital status	No	No	No	Yes

The sample is defined as in column 6 in Table 1. All adjusted models include quadratics in age, age at 1st marriage, age at 1st birth and years of education, interactions between education, age, age at 1st marriage and age at 1st birth as well as unrestricted dummies for state of birth, state of residence and residence in a SMSA. Heteroskedasticity-robust standard errors are in parentheses. All dollar figures are in 2002 constant dollars.

Table 5: Estimates of Women Economic and Labor Supply Models, by Age of Firstborn

Dependent variable:	(1) Entire Sample	(2) Firstborn Child <12	(3) Firstborn Child 12+
<u>**OLS**</u>			
Equivalized Household Income	-53.1	-48.5	-61.0
	(0.6)	(0.8)	(1.1)
Poverty	0.119	0.130	0.098
	(0.001)	(0.002)	(0.002)
Non-Woman Income	-20,856.1	-18,887.7	-24,664.3
	(104.0)	(124.1)	(187.5)
Woman's Income	8,680.8	8,275.0	9,540.6
	(50.1)	(58.0)	(94.5)
Wages	5,779.6	5,686.5	6,036.5
	(46.1)	(54.4)	(85.0)
Working for Pay	0.188	0.215	0.141
	(0.002)	(0.002)	(0.003)
Weeks per Year	10.163	11.046	8.641
	(0.081)	(0.100)	(0.139)
Hours per Week	9.663	10.564	8.152
	(0.067)	(0.083)	(0.113)
<u>**TSLS**</u>			
Equivalized Household Income	136.0	209.3	49.6
	(67.4)	(113.8)	(77.2)
Poverty	0.087	0.019	0.160
	(0.093)	(0.155)	(0.105)
Non-Woman Income	5,161.3	26,981.5	-19,491.3
	(10,147.6)	(17,840.7)	(12,181.7)
Woman's Income	13,041.5	3,830.7	22,931.0
	(4,487.8)	(6,888.6)	(6,209.6)
Wages	11,377.5	3,789.8	19,386.9
	(4,214.8)	(6,406.0)	(5,780.1)
Working for Pay	0.148	-0.101	0.415
	(0.179)	(0.302)	(0.207)
Weeks per Year	24.492	12.122	38.429
	(8.450)	(13.083)	(11.146)
Hours per Week	9.626	-2.030	22.559
	(6.784)	(11.352)	(8.251)
F-Statistic from First Stage	46.1	18.9	29.7
Sample Size	465,595	327,371	138,224
<u>For Ever-Divorced Women</u>			
% Living with Parent(s)	7.7	9.7	4.1
% Currently Married	44.1	40.9	49.9

The sample is defined as in column 6 in Table 1. All adjusted models include quadratics in age, age at 1st marriage, age at 1st birth and years of education, interactions between education, age, age at 1st marriage and age at 1st birth as well as unrestricted dummies for state of birth, state of residence and residence in a SMSA. Heteroskedasticity-robust standard errors are in parentheses. All dollar figures are in 2002 constant dollars.