Sex Preferences, Marital Dissolution, and the Economic Status of Women

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ABSTRACT

The rise in the divorce rate over the past 40 years is one of the fundamental changes in American society. A substantial number of women and children now spend some fraction of their life in single female-headed households, leading many to be concerned about their economic circumstances. Estimating the cause-to-effect relationship between marital dissolution and female economic status is complicated because the same factors that increase marital instability also may affect the economic status and labor market outcomes 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 continue following the birth of girls as opposed to boys. Our IV results cast doubt on the widely held view that divorce causes large declines in economic status for women. Once the negative selection into divorce is accounted for, our results show that, on average, ever-divorced women live in households with more income per person than never-divorced women.

I. Introduction

According to the Center for Disease Control (July 2002), 33 percent of first marriages now end in separation or divorce within ten years. The rising incidence of marital dissolution[1] has received substantial attention among social scientists and

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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.

policymakers. A large body of research has documented the lower economic status of unmarried women relative to married women and the decline in the economic status for women who experience marital dissolution (see Smock, Manning, and Gupta 1999 for a review of the literature). For example, Duncan and Hoffman (1988) find that the average white woman's family income falls by 30 percent in the year following a separation/divorce, and that even five years later the average woman has still only recovered to 88 percent of her preseparation family income. 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, as Smock, Manning, and Gupta (1999) point out, it is uncertain whether the negative association between divorce and the economic well-being of women represents a cause-to-effect relationship. First, divorce is more common within less educated and economically disadvantaged groups. Secondly, women who divorce also likely differ in ways that are not easily measured (see Holden and Smock 1991). Taken together, these suggest that the same factors that increase the probability of divorce also may be detrimental to the economic and labor market status of women. A limited number of studies have tried to address the nonrandom selection into divorce using parametric selection models and various exclusion restrictions (see Smock, Manning, and Gupta 1999 and the references therein). However, as is often the case with such approaches the results tend to be sensitive to the selected exclusion restrictions and the exclusions themselves are sometimes difficult to defend (parental education and employment status, for example). A convincing assessment of the impact of marital instability on the economic status of women therefore requires a credibly 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 (see Fuchs 1988 and Blau 1998 for overviews). 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 composed of never-married and divorced mothers. Because 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 to improve the efficiency of these programs. 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 dissolution 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 divorce. This exploits the fact that divorces are more likely following the birth of daughters than sons (Morgan, Lye, and Condran 1988; Teachman and Schollaert 1989; Katzev, Warner, and Acock 1994).[2] Using data from the 1980

2. Lundberg and Rose (2003) and Dahl and Moretti (2004) similarly find that the birth of a son out of wedlock increases the probability that the mother marries the father of the child.

Census, we find that the rate of marital dissolution is 4 percent 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 her first marriage.[3] Dahl and Moretti (2004) and Ananat and Michaels (2004) report similar estimates for the relationship between the sex of the firstborn child and the probability of marital dissolution using the 1980 Census.[4]

We implement this instrumental variables strategy using data from the 1980 U.S. Census of Population. As was previously observed, cross-sectional OLS comparisons of standardized (person-adjusted) household income and poverty rates among ever- and never-divorced women indicate a substantial economic disadvantage for ever-divorced women. In contrast, using firstborn sex as an exogenous determinant of marital dissolution, we find that ever-divorced women have significantly higher levels of standardized household income. However, the imprecision of IV estimates for poverty status limit our ability to make a definitive statement about the causal impact of marital instability on female poverty. Combined, these results suggest that OLS comparisons may overstate the detrimental effect of divorce on the economic status of women.[5]

Based on these findings, we investigate two channels through which divorced women might possibly improve their economic position: Labor market attachment and access to other sources of income. First, our IV results indicate that ever-divorced mothers earn approximately $5,000 more (in 1979 dollars)[6] per year more than their never-divorced counterparts. We show that this earnings advantage is due to greater labor-supply intensity (weeks and 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.

Divorced women also can improve their economic situation by obtaining more nonlabor income, welfare payments, and custody transfers, or cohabitating or remarrying. We investigate this possibility by comparing nonwoman income (total household income not generated by the woman) and nonwage income (a woman's total personal income minus her annual earnings) across ever- and never-divorced women. Again, simple cross-sectional comparisons indicate substantially lower nonwoman and nonwage income for ever-divorced women, while our IV estimates do not allow us to distinguish between zero and small negative effects. We also find that some divorced women move in with their parents, which increases their nonwoman income and compensates for the loss of their husband's income.

The remainder of the paper is as follows. Section II describes the data. Section III documents the relationship between the sex of the firstborn child and marital

3. In contrast, the sex of subsequent children may be endogenous due to nonrandom fertility, this point is also made by Rose (2000).

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 in the population of women with at least two children.

5. While doing the final revision of this paper we became aware of a study by Ananat and Michaels (2004) using a similar research design and 1980 Census data. For the variables common to both studies the results are similar. However, Ananat and Michaels (2004) differs from our analysis in terms of focus: they focus on the impact of divorces for the distribution of income while we focus on average income and labor-supply effects.

6. While the Census date is April 1, 1980, the reported income and earnings figures are for 1979.

dissolution. Section IV characterizes the population of ever-divorced women and demonstrates the random assignment of firstborn sex. Section V analyzes the relationship between marital dissolution and the economic status of women. The last section concludes.

II. Data

A. Identifying Ever-Married Mothers

The data for this study are drawn from the 1980 U.S. Census Public-Use Micro Samples.[7] 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 nationally representative data source of sufficient size allowing this kind of analysis. The 1990 and 2000 Censuses do not contain marital history information (only current marital status is reported), which makes the dating of the first marriage relative to the first birth impossible. Because our empirical strategy relies on identifying the sex of the first child born during the first marriage, we therefore cannot use data from the more recent Censuses. In contrast, while the 1960 and 1970 Censuses provide marital and fertility history information, the rate of marital dissolution was too low in the 1950s and 1960s to motivate an empirical investigation.

Table 1 reports the summary statistics and illustrates the construction of the sample used in the analysis. Our sample includes U.S.-born white women aged 21–40 who have been married at least once and who have had at least one child.[8] Throughout the analysis we focus on women whose first marriage began when they were 17–26 years of age. This includes more than 90 percent of all first marriages. The summary statistics for ever-married women who have had at least one child are reported in the first column of Table 1. As indicated in Column 1, this group had an average of 2.2 children and 26 percent of these women's first marriages have ended.

The Census does not provide any information about children who have moved out of the household. We therefore follow Angrist and Evans (1998) and limit the sample to women whose children all reside in her household. This restriction allows us to ascertain the sex of the oldest child by matching them to their mother within the household using the household relationship information in the Census.[9] This feature

7. The data were accessed from the Integrated Public Use Microdata Series Version 3.0 (Ruggles and Sobek 2003). 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.

8. We focus on white women because the high rate of out-of-wedlock births among black women renders a highly selected sample. Based on women aged 21–40 in the 1980 Census, 5 percent of births to white women occurred before their first marriage, while 46 percent of births to black women were before their first marriage.

9. To avoid confounding firstborn sex and family size we also exclude women with firstborn twins. Since the occurrence of firstborn twins is so small in the 1980 Census their inclusion or exclusion is immaterial.

of the Census introduces a missing-data problem: The characteristics of children who reside outside of the mother's household are not observed. This could potentially lead to biased estimates 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. However, there is little evidence to suggest that this is a problem. Using data from the 1980, 1985, and 1990 June CPSs we can analyze the patterns of living arrangements for children from disrupted families.[10] The June CPS contains complete marital and fertility histories, including information on the living arrangements of all children for each mother. We can therefore calculate the percentage of firstborn girls for the entire sample of ever-divorced women in the June CPS, including women with at least some children residing outside of her household Ĕ the group that is not observed in the Census sample. In the CPS sample, 50.2 percent of firstborn children are girls compared to 49.8 percent in the 1980 Census sample of ever-divorced women whose children all reside in her household. Therefore, firstborn girls are over-represented in our Census sample of ever-divorced mothers by at most 0.4 percentage points. Further, in the 1980 June CPS, the point estimate for the effect of firstborn girls on marital dissolution is 0.008, which is identical to the estimate we obtain in the 1980 Census (see Table 2).

For similar reasons we further exclude women whose oldest child is 18 or older. This restriction is necessary because youth are progressively more likely to leave their mother's home as they age. More importantly, the process of moving out of the mother's household may be nonrandom. In particular, girls tend to leave home at younger ages. In the 1980 Census sample of mothers whose children all reside in her household and whose oldest child is aged 18 or older, only 44 percent of firstborn children are girls, compared to 49 percent in the population at large.

The summary statistics for the subsample of ever-married women whose children all live in her household and whose oldest child is less than 18 years old are reported in Column 2 of Table 1. For this subsample, 21 percent of first marriages have ended in divorce. The lower incidence of marital instability among this group reflects two facts. First, this group is 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). Secondly, excluding women with children older than the age of 17 eliminates older women who have had a longer period of time over which to experience marital dissolution.

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. This time frame restriction limits the probability of adoption, which takes time, and reduces the probability that the oldest child in the household is a stepchild given the limited scope for multiple relationships in a short period.[11] Combined, these restrictions allow us to identify the sex of the firstborn child within

10. For comparability the sample definitions are identical to the restrictions placed on the Census sample.
11. All results are similar if this restriction is removed and the sample reported in the second column of Table 1 is used instead of the sample reported in Column 3.

Table 1
Descriptive Statistics

	(1) Ever-Married with Children	(2) All Children Live in Household	(3) 1st Child Born Within 5 Years of 1st Marriage
Marital History			
First marriage ended	0.26	0.21	0.20
	(0.44)	(0.41)	(0.40)
Age at first marriage	19.93	20.10	20.03
	(2.15)	(2.14)	(2.13)
Fertility			
Firstborn girl	—	0.49	0.49
		(0.50)	(0.50)
Children ever-born	2.18	2.03	2.08
	(1.08)	(0.93)	(0.94)
Age at first birth	—	22.62	22.17
		(3.21)	(2.68)
Socioeconomic Characteristics			
Age	31.38	30.61	30.53
	(5.15)	(4.84)	(4.89)

Years of education	12.66	12.80	12.74
	(2.10)	(2.09)	(2.02)
Urban	0.64	0.65	0.64
	(0.48)	(0.48)	(0.48)
Economic Status			
Standardized household income	9,858.3	9,786.7	9,623.3
	(5,600.4)	(5,478.3)	(5,339.8)
Poverty	0.07	0.07	0.07
	(0.26)	(0.25)	(0.25)
Nonwoman income	18,507.7	18,476.9	18,408.1
	(12,954.5)	(12,697.0)	(12,655.8)
Woman's income	4,640.5	4,417.7	4,322.8
	(5,947.0)	(5,846.9)	(5,743.4)
Annual earnings	4,044.3	3,863.1	3,776.9
	(5,433.0)	(5,331.4)	(5,227.4)
Sample size	662,204	535,887	465,595

Notes: The baseline sample includes U.S. born white women who are currently aged 21–40 who have been married at least once and had at least one child. Ever-married is defined as the first marriage beginning when the respondent was 17–26 years old. Standard deviation in parentheses. Columns progressively restrict the sample as defined in the column header.

the first marriage. The summary statistics for this final sample of 465,595 women are reported in Column 3. The similarity of the entries in Columns 2 and 3 indicates that the last restriction is innocuous. Except where noted otherwise, the empirical analysis in the remainder of the paper uses the sample described in Column 3. For expository ease, we refer to the final sample as "ever-married mothers." As indicated in Column 3, the divorce rate among ever-married mothers is 20 percent; average fertility is 2.1 children, and the average age at first birth 22.2.

B. Measuring the Economic Status of Women

There is no single widely accepted measure of "economic well-being." Reflecting this fact, we analyze several variables corresponding to different components of economic status. We begin by analyzing household income, which we standardize for the number of adults and minors younger than the age of 18 present in the household. Following the Census Bureau (2002) guidelines, household income is normalized by $(A + PK)^F$, where children (K) are defined as some proportion (P) of an adult (A) and F is the scale economy for adult-equivalent household size. Again following the recommendations of Census Bureau, we set F and P equal to 0.7.[12] All results are similar if household income is normalized by the household composition specific poverty line[13] (the poverty line for a household with a specific number of adults and minors), and are available from the authors upon request. We also consider poverty (=1 if the household's total income places them below the poverty line), nonwoman income (total household income minus the total income of the woman), personal income (the woman's total income), and the woman's annual earnings. 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.[14]

The average economic status of the women in our sample is reported in the bottom panel of Table 1. The average level of standardized household income among ever-married mothers is $9,623 (in 1979 dollars), and 7 percent of these women live in households that are below the poverty line. The Census Bureau (2003) similarly reports that the poverty rate for white families with children younger than the age of 18 was 8 percent in 1980. As we use the household as the unit for calculating poverty, our slightly lower poverty rate appears consistent with the Census Bureau figure. Further decomposing female economic status into its components reveals that nonwoman income, household income minus all forms of income not generated by the woman, constitutes over 80 percent of household income for women, and that the remainder is almost entirely female labor market earnings. Although, the woman's nonlabor sources of income, such as welfare payments, child support, and alimony do not appear to be important income sources for the average ever-married mother, they do differ across ever- and never-divorced groups—we return to this issue in Section V.

12. The Census Bureau recommends setting $P = 0.7$ and $0.65 \leq F \leq 0.75$.

13. This accounts for the fact that larger households with the same household income are worse off. See *http://www.ipums.umn.edu/usa/volii/1990Poverty.htm* for family poverty definitions, which we translate to households.

14. All results are very similar using family measures and are available from the authors upon request.

III. The Relationship Between Firstborn Sex and Marital Dissolution

The fact that the sex composition of offspring, in families with at least two children, affects subsequent fertility is often interpreted as evidence of parental preference for mixed-sex composition among their offspring (examples include Ben-Porath and Welch 1976; Leung 1991; Angrist and Evans 1998). Others have suggested that couples may have a preference for firstborn male children (Williamson 1976). However, parental sex preferences also may influence marital stability. For example, it has been documented that marriages are less likely to continue after the birth of daughters than sons (Morgan, Lye, and Condran 1988; Teachman and Schollaert 1989; Katzev, Warner, and Acock 1994; Dahl and Moretti 2004; Ananat and Michaels 2004). 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.[15] We also explore the potential sources of heterogeneity in the relationship between marital instability and firstborn sex. Column 1 reports the unadjusted differences and associated F-statistics testing the null hypothesis that firstborn sex has no effect on marital instability and Column 2 reports the regression-adjusted differences and associated F-statistics.[16] Finally Column 3 displays the number of observations.

Panel A reports the overall effect of firstborn sex on the probability of marital disruption. The point estimate is 0.008 (standard 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 percent higher divorce rate for women with firstborn girls relative to firstborn boys given an average divorce rate of 20 percent for 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 random. The large F-statistics (50.2 and 46.1 for the unadjusted and adjusted models, respectively) confirm the importance of the sex of the firstborn child on marital instability and are well above the rule-of-thumb values suggested by Bound, Jaeger, and Baker (1995) and Staiger and Stock (1997) in their studies of weak instrumental variables.

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,

15. All models are estimated as linear probability models (OLS). For completeness, the full set of parameters for the first-stage models are reported in Appendix Table 2.

16. 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. As a sensitivity check, we also ran all models excluding age at first marriage and age at first birth, given their potential endogeneity. As the results are essentially identical when these variables are excluded we do not report the estimates in the paper, however, the results are available from the authors upon request.

Table 2
Effect of Firstborn Sex on the Probability of Marital Instability

	(1) Unadjusted		(2) Regression Adjusted		(3)
Dep = First Marriage Ended	Coefficient	*F*-Statistic	Coefficient	*F*-Statistic	Observations
(A) Overall Effect					
Firstborn girl	0.008	50.2	0.008	46.1	465,595
(B) By Education Level					
<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
(C) By Age at First Marriage					
<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
(D) By Age at First Birth					
<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

Notes: The sample is defined as in Column 3 in Table 1. The adjusted models include quadratics in age, age at first marriage, age at first birth and years of education, interactions between education, age, age at first marriage and age at first birth, as well as unrestricted dummies for state of birth, state of residence and residence in a SMSA. The complete first stage results are reported in Appendix Table 1. Estimated by OLS.

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 firstborn 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 22nd 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.

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 percent higher divorce rates for women with firstborn girls, relative to the average divorce rates for these years. These results are reported in Appendix Table 1 (also see Dahl and Moretti 2004).

IV. The Random Assignment of Firstborn Sex and the Characteristics of Ever-Divorced Women

In the neoclassical theory of marriage (Becker 1973 and 1974), marital gains are derived from specialization within the household and hence depend on the woman's potential earnings capacity relative to her husband's, which is determined by the marital matching process (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 of 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. All differences reported in Column 3 are statistically significant at the 5 percent level. 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.

In contrast to the significant 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 women with firstborn girls and boys,

Table 3
Differences 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

Notes: The sample is defined as in Column 3 in Table 1. The standard deviations reported in parentheses except in Columns 3 and 6 where the entries in parentheses are heteroskedasticity-robust standard errors. Differences estimated by OLS.

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. 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. 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.

V. The Effect of Marital Dissolution on the Economic Status of Women

A. OLS Estimates

We begin the empirical analysis by using OLS regressions to estimate the relationship between female economic status and marital instability.[17] Let Y_i denote a measure of economic status for woman i:

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

As previously stated, we focus on five indicators of economic status: standardized household income, poverty, nonwoman income, personal income, and annual earnings. In addition, we also analyze the labor-supply determinants of labor market earnings, including 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, years of education, a dummy for SMSA status, unrestricted state of birth and state of residence dummies, and interactions between years of education and all of 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 of the effects in Column 1 are very precisely estimated. Standardized household income is \$1,573 lower for ever-divorced mothers and the poverty rate is 12 percentage points higher.[18] This is entirely explained by the large reduction in nonwoman income following divorce. As Column 1 shows, ever-divorced women have higher average levels of 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 earnings arise from the stronger labor market attachment of ever-divorced women. Taken together, these estimates are qualitatively similar to those in other studies (Johnson and Skinner 1986; Duncan and Hoffman 1985 and 1989; Burkhauser et al. 1991; Smock, Manning, and Gupta 1999).

17. For the binary outcomes (like poverty status and employment), the OLS models correspond to linear probability models. In all cases the probit estimates of the marginal effects are nearly identical.
18. Similar OLS and IV results are obtained when poverty is replaced by public assistance receipt.

Table 4
The Effect of Divorce on Female Economic Status and Labor Supply

Dependent variable	(1) OLS	(2) WALD	(3) TSLS	(4) TSLS
Standardized household income	−1,572.6	4,142.1	3,921.2	6,301.9
	(19.5)	(2,048.5)	(2,018.7)	(3,395.3)
Poverty	0.119	0.068	0.087	0.143
	(0.001)	(0.088)	(0.093)	(0.146)
Nonwoman income	−9,553.9	2,966.8	2,364.3	2,217.5
	(47.6)	(4,617.3)	(4,648.5)	(7,403.2)
Woman's income	3,976.5	6,322.4	5,974.1	9,606.0
	(22.9)	(1,980.7)	(2,055.8)	(3,291.7)
Annual earnings	2,647.6	5,495.0	5,211.9	8,696.1
	(21.1)	(1,854.7)	(1,930.7)	(3,130.6)
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)
Includes fertility and current marital status	No	No	No	Yes

Notes: Sample as defined in Column 3 in Table 1. All adjusted models include quadratics in age, age at first marriage, age at first birth, and years of education, interactions between education, age, age at first marriage and age at first birth and indicators for state of birth, state of residence and residence in a SMSA. Heteroskedasticity-robust standard errors are in parentheses.

B. TSLS Estimates

Although 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 significant differences in observable characteristics across never- and ever-divorced women. For example, ever-divorced women are 0.9 years younger at the time of their first marriage, 1.2 years younger at the time of their first birth, and have 0.4 years less education than never-divorced women. These differences in observable determinants of divorce suggest that the populations of never- and ever-divorced women also may differ along unobservable dimensions (Altonji, Elder, and Taber 2000). Because the observable characteristics of divorced mothers are typically associated with lower economic status (less education and younger ages at first birth for example), it seems plausible that ever-divorced women have worse unobserved determinants of economic status and labor supply. This would tend to bias the

OLS estimates downward, even after conditioning on the observable determinants of economic status. To illustrate, consider the mean difference for Y_i in Equation 1 by divorce status:

(2) $$E[Y_i | D_i = 1, X_i] - E[Y_i | D_i = 0, X_i] = \beta + E[\varepsilon_i | D_i = 1, X_i] - E[\varepsilon_i | D_i = 0, X_i]$$

As Equation 2 indicates, unless $E[\varepsilon_i|D_i = 1,X_i] = E[\varepsilon_i|D_i = 0,X_i]$, OLS estimates of β will be biased. Moreover, if there is negative selection on unobservables into divorce ($E[\varepsilon_i|D_i = 1,X_i] < E[\varepsilon_i|D_i = 0,X_i]$), the OLS estimates will be biased downward.

We propose an instrumental variables solution to this problem using the sex of the firstborn child as an exogenous determinant of divorce. The sex of the firstborn child is a valid instrument for studying the impact of divorce on female economic status if two requirements are satisfied. First, the sex of the firstborn child must be correlated with the probability of marital instability. Table 2 clearly shows this to be the case; firstborn sex is an important determinant of marital disruption with an F-statistic of 46.1. Second, the sex of the firstborn child must be uncorrelated with the unobserved determinants of female economic status. In other words, the second assumption presumes that firstborn sex influences economic status only through its affect on marital stability, so that firstborn sex can be rightfully excluded from models like Equation 1. This assumption is violated if, for example, offspring sex changes the behavior of either the mother or father in ways that directly impact the economic status and/or labor supply of women. However, without additional exclusion restrictions this assumption is not testable.

Instead, we present two pieces of evidence suggesting that firstborn sex is a credible instrument to study the effect of divorce on the economic status of women. First, we reexamine the results of Lundberg and Rose (2002). Using a fixed effects approach, they find that men born during the 1940s and 1950s in the PSID work approximately one more hour per week after the arrival of a son. In contrast, they find no such effect on the labor supply of mothers and no wage effect for either mothers or fathers. Combined, these provide little evidence of a direct effect of child sex on female labor market behavior or wages and imply at most a very limited impact on female economic status through male labor market behavior.[19] Second, as suggested by Angrist, Imbens, and Rubin (1996), and Angrist and Krueger (1999) we test for any association between firstborn sex and predetermined predictors of divorce (age, age at first marriage, and education). A credible instrument should be uncorrelated with observable variables that are determined *before* divorce can take place. The results reported in Columns 4–6 in Table 3 confirm this, thus providing no evidence against the null hypothesis that firstborn sex is unrelated to the unobserved determinants of female economic status.

To exploit the randomness embodied in the sex of the firstborn child, we use an indicator for firstborn girls as an instrument for divorce. More specifically, we estimate the parameters of Equation 1 using TSLS based on the following the first-stage equation for divorce:

(3) $$D_i = \pi_0 + \pi_1 G_i + X_i \pi_2 + \nu_i .$$

where G_i takes a value of one if the firstborn child is a girl and zero if it is a boy. Estimates of π_1 have already been reported in Table 2. Column 2 in Table 4 reports

19. The increase in hours worked associated with a firstborn boy results in an additional \$500 in male earnings, which corresponds to a 2 percent increase in household income for never-divorced mothers.

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 similarly reports the TSLS estimates when the control variables are included in the models. While the model in Equation 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; Angrist 2003). 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 her firstborn child.

The entries in Columns 2 and 3 point to two main findings: First, to the extent that economic status is adequately measured by standardized household income, there is no evidence of a negative impact of divorce on economic status. Second, the higher levels of personal income and annual earnings for ever-divorced women are attributable to greater labor-supply intensity rather than labor market participation. These conclusions are supported by two sets of results.

The TSLS estimates for standardized household income are higher for ever-divorced women. In all cases, this result is both statistically significant at the conventional level and the Hausman tests reject the null hypothesis that the TSLS and OLS estimates are the same (except for sampling errors) at the 5 percent level. At the same time, the TSLS estimates indicate that personal income and annual earnings are \$5,974 and \$5,212 higher on average among ever-divorced women (both differences are statistically significant at the 5 percent level). However, the evidence on poverty and nonwoman income is weaker: For both outcomes, the TSLS estimates are not significantly different from zero. The absence of a significant impact of divorce on poverty in our specifications may stem from the coarseness of our poverty measure. When Ananat and Michaels (2004) look at the probability of falling below a variety of family income cutoffs, some below the poverty line and some slightly above, they find that ever-divorced women are more likely to fall into low-income categories, even though there is no impact on average income.

What explains these results? As shown below, once negative selection is adjusted for, women respond to divorce by increasing labor-supply intensity. Another possibility is that divorced mothers move in with their parents in the short run and remarry in the long run, both of which increase total household income.

The higher earnings of ever-divorced women are mostly attributable to greater labor-supply intensity, and not participation. The TSLS estimates in Column 3 reveal that the average ever-divorced woman works 24 more weeks per year than the average never-divorced woman, 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 and the OLS estimates reported in Column 1 (those are 0.19 (employment rate) and 9.7 (hours per week)). However, as we will show in Section VC, the average effects hide important labor-supply differences across divorced women with younger and older children.

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 3, which shows that the sex of the firstborn child 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 nonrandom selection into remarriage.[20] The balancing property of firstborn sex documented in Table 3 is therefore lost once we condition on the 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 for all of the 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.

One related issue is the potential presence of omitted-variables bias in the TSLS estimates. In our application this bias is inversely proportional to the first-stage effect of firstborn sex on divorce. Despite statistical significance that satisfies the rule-of-thumb criterions for weak instruments (see, for example, Bound, Jaeger, and Baker 1995; Staiger and Stock 1997), the first-stage effects are small in magnitude and thus omitted-variables bias may confound some of our estimates. Since the first-stage effects are positive the sign of the omitted-variables bias is determined by the sign of $E[\varepsilon_i|Z_i = 1,X_i] - E[\varepsilon_i|Z_i = 0,X_i]$. If the mean of the unobservables is higher in the population with firstborn boys, then our TSLS estimates will be biased downward. Only in the case where the unobservables have a higher mean in the firstborn girl population are the TSLS estimates possibly biased upward, and Lundberg and Rose (2002) provides little evidence supporting this possibility.

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 child for several reasons.[21] First, independent 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 young children (relative those with older children). Finally, since the husband's income grows over time (because of the increasing age-earnings profile) the decline in nonwoman 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 are also somewhat more likely to move in with their parents following a divorce, which could alleviate the loss of their husband's income.

We explore these possibilities in Table 5 by estimating the models on samples of women whose firstborn is younger or older than 12 years of age. For comparison,

20. For example, ever-divorced mothers who remarry are older and more educated than those who do not.

21. We elected to condition on firstborn age because women with firstborn 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.

Column 1 reports the results from the full sample (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 women illustrated in Table 4.

In the bottom panel of Table 5 we report TSLS 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 children are aged 12 and 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, we note that the positive TSLS effect of divorce on personal income and annual earnings reported in Table 4 is concentrated in the group of women whose oldest child is at least 12 years old. The personal income and earnings differentials are \$10,504 and \$8,881 respectively, and are precisely estimated. Among women with younger firstborn children the TSLS estimates on personal income and annual earnings are small and not statistically significant. This reflects the fact that the labor-supply intensity of divorced mothers increases with firstborn 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). Using a similarly aged[22] panel of women from the PSID, Johnson and Skinner (1986) also find that hours of work rise by 300–650 per annum after separation, depending on the exact year before and after separation that you compare. The differential impact of divorce on labor supply across women with older and younger children is not surprising given the differential child-care costs across these groups. It is also worth highlighting the fact that these results imply that ever-divorced women, at least those with older children, enjoy less leisure time than never-divorced women with similarly aged children.

Next we turn to standardized household income. Again, the TSLS estimates differ substantially across the two subsamples. The estimated TSLS differences in standardized household income across divorce status are \$6,066 (standard error = 3,374) in the group with young children and \$1,357 (standard error = 2,362) in the group with older children. Although, the point estimate for women whose oldest child is 12 years old or older is relatively imprecise, it is clear that the estimated higher standardized household income for the entire sample of ever-divorced is driven by women with young children. This partly reflects the fact that 10 percent of divorced mothers with young child were living with their parents in 1980. Because this group of women is relatively young, it is plausible that the parents of divorced mothers have higher earnings than the husbands of mothers who have remain married, which would tend to increase nonwoman income among young divorced mothers. Second, among mothers with older children the negative TSLS estimate on nonwoman income can be attributed to the increasing age-earnings profile of husband of never-divorced women, which makes the costs of divorce higher in terms of lost nonwoman income.

22. Compared to our sample of women whose oldest child is at least 12 years old.

Table 5
The Effect of Divorce on Female Economic Status and Labor Supply, by Firstborn Age

Dependent variable:	(1) Entire Sample	(2) Oldest Child <12	(3) Oldest Child 12+
OLS			
Standardized household income	–1,572.6	–1,400.0	–1,885.4
	(19.5)	(23.4)	(34.9)
Poverty	0.119	0.130	0.098
	(0.001)	(0.002)	(0.002)
Nonwoman income	–9,553.9	–8,652.2	–11,298.3
	(47.6)	(56.9)	(85.9)
Woman's income	3,976.5	3,790.7	4,370.4
	(22.9)	(26.6)	(43.3)
Annual earnings	2,647.6	2,604.9	2,765.3
	(21.1)	(24.9)	(38.9)
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)
TSLS			
Standardized household income	3,921.2	6,066.4	1,357.0
	(2,018.7)	(3,374.0)	(2,361.5)
Poverty	0.087	0.019	0.160
	(0.093)	(0.155)	(0.105)
Nonwoman income	2,364.3	12,359.8	–8,928.7
	(4,648.5)	(8,172.6)	(5,580.3)
Woman's income	5,974.1	1,754.8	10,504.3
	(2,055.8)	(3,155.6)	(2,844.5)
Annual earnings	5,211.9	1,736.1	8,880.9
	(1,930.7)	(2,934.5)	(2,647.8)
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

Notes: Sample as defined in Column 3 in Table 1. All adjusted models include quadratics in age, age at first marriage, age at first birth, and years of education, interactions between education, age, age at first marriage and age at first birth and indicators for state of birth, state of residence, and residence in a SMSA. Heteroskedasticity-robust standard errors are in parentheses.

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

In this paper we present evidence based on an instrumental variable 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 percent more likely to experience marital dissolution than families with firstborn boys. Because 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 nonrandom 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 substantial economic disadvantage for divorced women. In particular, divorce is associated with higher poverty rates, lower standardized household income, and lower nonwoman income (total household income not generated by the woman). Similar findings have previously been interpreted by some analysts as indicating that marital status is an important causal determinant economic well-being for women. However, the evidence from our IV models brings this conclusion into question. Once the negative selection of women into divorce is accounted for, ever-divorced mothers have substantially higher levels of personal income and annual earnings than never-divorced mothers. These effects are precisely estimated. We also show that this advantage is concentrated among mothers with older children. The better labor market performance of divorced women is an important factor in improving their economic position. Further, our IV estimates point to higher standardized household income for ever-divorced women. This result is partly explained by the fact that some divorced mothers with young children tend move in with their parents, thereby raising their standard of living. As such, our IV evidence for standardized household income is not consistent with the contention that marital dissolution causes large declines in economic status. Tempering these results, our estimates for poverty incidence are always too imprecise to make conclusive statements about the cause-to-effect relationship between marital instability and poverty. Together these results cast doubt on the contention that divorce necessarily causes a reduction in the economic status of women.

Finally, we show that the higher labor market earnings of ever-divorced women are almost entirely attributable to increased labor-supply intensity, relative to continuously married women. Our IV estimates indicate that divorce is associated with more hours and weeks worked, and thus higher labor market earnings. At the same time, we find little difference in the employment rates of ever- and never-divorced women. This evidence leads us to conclude that marital dissolution affects labor-supply intensity but not the decision to participate in the labor market.

Appendix Table A1

Marital Instability by the Sex of the Firstborn Child

Dep = First Marriage Ended	Firstborn Girl	Firstborn Boy	Unadjusted Difference	*F*-Statistic	Reg-Adjusted Difference	*F*-Statistic
1960 (n = 85,508)	0.090	0.083	0.006	10.6	0.004	4.5
	(0.286)	(0.276)	(0.002)		(0.002)	
1970 (n = 256,681)	0.123	0.118	0.005	13.1	0.004	11.5
	(0.328)	(0.323)	(0.001)		(0.001)	
1980 (n = 465,595)	0.203	0.195	0.008	50.2	0.008	46.1
	(0.402)	(0.396)	(0.001)		(0.001)	

Notes: The sample is defined as in Column 3 in Table 1. Standard deviations in parentheses except in the unadjusted difference column where heteroskedasticity-robust standard errors are in parentheses.

Appendix Table A2
First Stage Results

	Overall	By Education Level				By Age at First Marriage		By Age at First Birth	
Dep = First Marriage Ended		<12	12	13–15	16+	<20	20+	<22	22+
Firstborn Girl	0.008	0.016	0.006	0.009	0.004	0.010	0.006	0.011	0.005
	(0.001)	(0.004)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.002)	(0.001)
(A) Age	0.052	0.059	0.072	0.068	0.008	0.070	0.050	0.083	0.046
	(0.001)	(0.005)	(0.002)	(0.005)	(0.007)	(0.003)	(0.002)	(0.003)	(0.002)
(B) Age at first birth	–0.060	–0.077	–0.096	–0.133	–0.084	–0.180	–0.099	–0.074	0.018
	(0.004)	(0.020)	(0.006)	(0.018)	(0.021)	(0.014)	(0.007)	(0.034)	(0.008)
(M) Age at first marriage	–0.141	–0.043	–0.108	–0.217	–0.144	0.132	–0.048	–0.266	–0.217
	(0.006)	(0.026)	(0.009)	(0.023)	(0.027)	(0.075)	(0.012)	(0.033)	(0.008)
(E) Years of education	0.023	–0.009		0.056	0.066	0.116	–0.012	0.102	–0.025
	(0.003)	(0.014)		(0.070)	(0.050)	(0.016)	(0.005)	(0.013)	(0.005)
A-Squared	–0.104	–0.094	–0.110	–0.111	–0.082	–0.142	–0.093	–0.165	–0.084
	(0.002)	(0.008)	(0.003)	(0.005)	(0.006)	(0.004)	(0.003)	(0.004)	(0.003)
E-Squared	0.101	0.058		0.171	0.032	0.204	0.060	0.182	0.070
	(0.008)	(0.032)		(0.255)	(0.138)	(0.016)	(0.009)	(0.015)	(0.009)
M-Squared	0.326	0.082	0.230	0.644	0.367	–0.332	0.110	0.743	0.476
	(0.016)	(0.057)	(0.216)	(0.035)	(0.041)	(0.210)	(0.028)	(0.091)	(0.019)
B-Squared	0.189	0.144	0.172	0.256	0.277	0.577	0.212	0.310	–0.028
	(0.010)	(0.039)	(0.014)	(0.023)	(0.025)	(0.035)	(0.015)	(0.088)	(0.017)
A*E	0.145	–0.002		0.069	0.318	0.166	0.113	0.169	0.105
	(0.006)	(0.024)		(0.032)	(0.037)	(0.014)	(0.007)	(0.014)	(0.007)
M*E	–0.069	–0.113		–0.530	–0.163	–0.407	–0.015	–0.238	–0.008
	(0.028)	(0.112)		(0.151)	(0.141)	(0.110)	(0.032)	(0.088)	(0.029)
B*E	–0.356	0.133		–0.071	–0.462	–0.653	–0.176	–0.715	–0.134
	(0.024)	(0.096)		(0.121)	(0.115)	(0.053)	(0.026)	(0.088)	(0.027)
Urban	0.041	0.049	0.041	0.037	0.027	0.058	0.023	0.058	0.024
	(0.001)	(0.004)	(0.002)	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)
F-statistic for firstborn girl	46.1	18.1	13.7	15.9	3.4	29.7	16.6	36.3	12.3
Sample size	465,595	51,981	247,053	102,126	64,435	216,822	248,773	207,584	258,011

Notes: Samples as defined in Table 2. All models also include indicators for state of birth and state of residence. Heteroskedasticity-robust standard in parentheses.

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