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Does Single Parenthood Increase the Probability of Teenage Promiscuity, Drug Use, and Crime? Evidence from Divorce Law Changes

Heather Antecol
Department of Economics
Claremont McKenna College
heather.antecol@mckenna.edu

Kelly Bedard
Department of Economics
University of California
Santa Barbara
kelly@econ.ucsb.edu

Eric Helland
Department of Economics
Claremont McKenna College
eric_helland@mckenna.edu

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Abstract

It has long since been established that children raised by single parents are more likely to become sexually active, commit illegal acts, and use illegal drugs at young ages. What has not been determined is whether or not there is a causal effect associated with the disintegration of the family. Would these children have been more likely to participate in 'deviant' behavior even if their family structure had remained intact? This study provides evidence in favor of a negative causal impact of single-parent status. Using state-level divorce law changes to instrument for years that the biological father lives in the household, we find that youth who spend part of their childhood/youth living in a household that does not include their biological father are more likely to smoke regularly, become sexually active, and be convicted of a crime.

Key Words: Family Structure, Marital Dissolution, Youth Outcomes
JEL Classifications: J12, J13

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1. Introduction

There is a large literature that seeks to measure the importance of family structure in determining child/youth outcomes. These studies show that children raised in single parent homes tend to perform more poorly in school, and are more likely to become sexually active, commit illegal acts, and use illegal drugs at young ages (examples include, Painter and Levine 2000, Comanor and Phillips 1998, Wu 1996, Garasky 1995, Manski et. al 1992, Astone and McLanahan 1991, Flewelling and Bauman 1990, and Matsueda and Heimer 1987).

In general, these studies use techniques, such as Ordinary Least Squares (OLS) or Probit/Logit analysis, to estimate single-equation models that seek to explain a child/youth outcome in terms of a number of individual characteristics, family characteristics, regional characteristics, and a measure of family structure, such as a dummy variable indicating whether the parents are divorced.[1,2,3] While these types of studies demonstrate the correlation between family structure and youth outcomes, the direction of causation is far from clear. For example, unobserved maternal ability may be associated with both poor parenting and family structure. Once the endogeneity of family structure is recognized, it is clearly inappropriate to estimate these types of models using estimation techniques, such as OLS or Probit/Logit analysis that do not account for endogeneity.[4]

[1] Johnson and Mazingo (2000) estimate the impact of parental divorce on later labor market and marriage outcomes for the children of divorce. While the Census data sets they use do allow them to construct instrument-like estimates of the impact of divorce, they are forced to use a two-sample instrumental variables (2SIV) estimator and are unable to control for important family background factors.

[2] Astone and McLanahan (1991) use bivariate probit analysis and Manski et al. (1992) use trivariate probit analysis to examine the effect of family structure on high school completion. Although these estimation techniques are more complex than the single-equation probit analysis, they are not useful in determining the causal relationship between family structure and youth outcomes because the results are sensitive to the exclusionary restrictions chosen.

[3] Corak (2001) uses adolescents who lose a parent to death as a benchmark for measuring the impact of divorce during adolescence on subsequent adult outcomes. He finds that children whose parents divorce during their adolescence tend to marry later and suffer greater marital instability, but he finds little impact on earnings.

[4] Painter and Levine (2000) point out that the absence of a clear understanding of causation is particularly problematic from a policy standpoint. If individual parental/socioeconomic characteristics are the driving force

The purpose of this paper is to add to the current debate about the impact of family structure on youth outcomes. In particular, we examine the impact of the length of time that children are exposed to their biological father on youth participation in smoking, drinking, sexual intercourse, drugs, and crime before the age of fifteen using both single-equation models that assume the exogeneity of family structure and simultaneous equation models that allow for the endogeneity of family structure.

We find that single- and simultaneous equation models generate quite different pictures of the importance of family structure. In the single-equation models, we find a moderate and statistically significant family structure effect on smoking regularly, sexual intercourse, and conviction before the age of fifteen. In the simultaneous equation models, the family structure effects are substantially larger. In particular, the single-equation estimates suggest that an additional year with the biological father decreases the probability of smoking regularly by 0.5 percentage points, engaging in sexual intercourse by 1.0 percentage point, and conviction by 0.1 percentage points. In contrast, the preferred (just identified) simultaneous equation estimates for one more year with the biological father results are 3.8, 3.9, and 0.5 percentage point decreases in the probability of smoking regularly, engaging in sexual intercourse, and conviction, respectively. However, the conviction coefficient is no longer statistically significant at conventional levels. Our results suggest that by ignoring simultaneity, previous studies substantially understated the effect of family structure.

The remainder of the paper is as follows. Section 2 describes the parental, youth, and divorce law data. Section 3 discusses the estimation and results from single-equation probit

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behind poor youth outcomes then policies attempting to keep families intact will have little effect on youth outcomes. On the other hand, if it is in the absence of the father in the household that leads to poor youth outcomes, then such policies might have a positive impact.

models (where family structure is treated as exogenous) and two-stage probit models (where family structure is treated as endogenous). Section 4 concludes.

2. Data

All youth, parental, and family data are drawn from the Geocode version of the National Longitudinal Survey of Youth (NLSY) and the NLSY Young Adult Supplement (NLSY-YAS). These data suit our purposes for a number of reasons. First, the NLSY-YAS allows us to include a wide range of youth outcomes, that is, participation in smoking, drinking, sexual intercourse, drugs, and crime before the age of fifteen. To the best of our knowledge no other study has included such a wide range of outcome measures. Secondly, the NLSY and the NLSY-YAS contain a broad range of control variables for youth as well as their parents, which is important as it allows us to identify a large set of pre- and post-divorce factors.[5] Thirdly, and most importantly, combining these data allows us to measure the length of time each youth lives with his/her biological father.

The sample is restricted to mothers and children residing with their mother during their entire first fifteen years of life. We restrict our attention to children living with their mother throughout their life because the small number of children raised by single fathers and alternate caregivers are too small to reliably analyze.

Since 1986 the children of NLSY women have been surveyed biannually. Child cognitive ability and development are assessed using tests and mothers are extensively surveyed

[5] While many studies include controls for family characteristics, the findings are mixed. Thomson et. al. (1994) find that family income explains more of the difference between children of intact families and children of single parent families than divorce itself. Furstenberg and Teitler (1994) show that controlling for pre-divorce family background factors, the quality of parental marital relations, child characteristics, and parent-child interaction greatly reduces the estimated impact of divorce on a child's well being. In contrast, Morrison and Cherlin (1995) find that controlling for pre-divorce factors has little impact on the estimated effect of divorce on young boys.

to establish the quality of the home environment. In 1994 the survey was extended to survey 'youth' aged fifteen and over directly. Each youth completes an interview focusing on education, employment, and family-related behavior as well as filling out a confidential questionnaire that focuses on substance use, sexual activity, and other such sensitive issues. In particular, youth are asked how old they were when they first smoked cigarettes and how often they have engaged in this behavior, began drinking alcohol at least once a month, engaged in sexual intercourse, used marijuana and how often they engaged in this behavior, and were convicted of a crime other than a minor traffic offense. This information is used to construct variables indicating whether or not the respondent participated in a specified 'deviant' behavior before the age of fifteen.[6] Behavior is measured at age fourteen to maintain a representative sample. In particular, older youth samples are less representative because they necessarily imply the over-sampling of individuals born to women who were very young at the point of childbirth.

Again to maintain the largest and most representative sample possible, the retrospective 'deviant' youth behavior reports for 1998 are used. A youth is only included if they are fifteen or older at the interview date so that behavior occurring up until the end of age fourteen is included. Table 1 reports the summary statistics for the sample.[7,8] Approximately 4 percent of youth are convicted of a crime before the age of fifteen, while 19 percent, 18 percent, 16 percent, and 15 percent become sexually active, smoke regularly, use marijuana regularly, and drink regularly, respectively.

[6] A respondent is coded as participating in smoking (or using marijuana) before the age of fifteen if they report participating for the first time before the age of fifteen and have smoked cigarettes (marijuana) on more than one hundred (ten) occasions.
[7] Sample size varies across dependent variables because of non-response. The sample sizes are 1258, 803, 1297, 861, 1282 for smoking, drinking, sex, marijuana use, and conviction samples, respectively. The summary statistics for the independent variables are based on the sex sample, however, similar results are found if we utilize any of the other dependent variable samples and are available from the authors upon request.
[8] 1998 sampling weights are used in all the tables.

The deviant behavior variables are linked to youth and parental control variables measured in the year in which the youth is fifteen years old. The youth's gender and birth-order are obtained from the NLSY-YAS. The sample is evenly split between male and female children, and approximately 63 percent of the sample are the first-born child. The mother's years of education, the number of siblings that the youth has, urban/rural residential location, the mother's census division of residence at age fourteen, and the youth's census division of residence at age fifteen are drawn from the NLSY.

Finally, and most importantly, combining the NLSY and the NLSY-YAS allows us to measure the length of time that each youth lives with his/her biological father (the maximum is 180 months - their entire life up to the age of fifteen). Although these data do not allow one to directly link youth to their biological father, we are able to link them through the mother's marital status. In particular, a man is considered to be the youth's father if he was either married to the mother at the point of birth or married her within 36 months of the youth's birth. In both cases, the father is assumed to be present in the household from the point of birth. If the father marries the mother at or before the youth's birth, the number of months that the father is present is measured by the number of months that the marriage lasts. If, on the other hand, the father marries the mother after the youth's birth, the number of months that the father is present is measured by the number of months that the marriage lasts plus the number of months from the youth's birth until the marriage began. In all other cases, we assume that the youth never lives with his/her biological father.

According to Table 2a, mothers who were unmarried at the time of the youth's birth but who married within the first 36 months of the youth's life (7 percent of the sample) are similar to women who married at or before the youth's birth (73 percent of the sample). In particular, both groups of women tend to remain married for longer periods of time compared to women whose first marriage (during the youth's life) occurs more than 36 months after the respondent's birth (10 percent of the sample). The mean marriage duration for women married at or before the time of the youth's birth and women married within 36 months of the respondent's birth are 145 and 124 months, respectively. In contrast, the average marriage of a woman whose first marriage (during the youth's life) occurs more than 36 months after respondent's birth is 77 months. Part of this difference is, of course, driven by the fact that for our purpose marriage duration is capped at 180 months. Given our inability to directly link youth and their fathers and the similarity of the marriage duration across mothers married at birth and married within 36 months, we believe that the more expansive father definition is descriptively more accurate. It should be noted, however, that our results are not sensitive to the father definition used; these issues are discussed in greater detail in Section 3.3.

One might also be concerned that it is the presence of a 'father figure' rather than the presence of a biological father that affects youth participation in deviant behavior.[9] To account for this we similarly measure the length of time that a child lives with a stepfather(s). Table 2b shows that 20 percent of youth whose mother married their first husband (the first from the viewpoint of the youth in question) at or before his/her birth have a stepfather(s) before the age of fifteen. Similarly, 26 percent of youth whose mothers married their first husband within 36 months of the youth's birth have a stepfather(s) before the age of fifteen. On average, the two aforementioned groups live with their first stepfather for 67 and 55 months, respectively. In contrast, 10 percent of youth have mothers who marry for the first time after they are 36 months

[9] In a similar vein, one might be concerned that changes in child support payments are also an important determinant of youth outcomes. While it is true that there were major changes to the Child Support Enforcement in 1984 and 1988, Case, Lin and McLanahan (2000) find that both the probability of receiving child support payments and the

of age and hence never live with their biological father. However, by definition these individuals do have a stepfather(s). On average, their first stepfather arrives when they are 85 months of age and remains in the household for 77 months. Given the relative stability of marriages to stepfathers, excluding time spent with these 'father figures' may bias the time spent with the father estimates downward by attributing good outcomes to youth defined as not having a 'father'. This is discussed in greater detail in section 3.3.

The Geocode version of the NLSY reports the mother's state of residence at age fourteen and in all survey years from 1979-98. Given this information we can construct a mother's exposure to different divorce laws, which we use to instrument for the length of time that the biological father is present (for a detailed discussion of the validity of liberalization of divorce laws as instruments see Section 3.2). The year in which each state enacted irretrievable breakdown (no-fault divorce) and no-fault based property division and alimony laws are listed in Appendix Table A1.[10] More specifically, we construct the following four measures for each mother: did your state of residence at age fourteen have a unilateral divorce law, years of exposure to unilateral divorce laws since age twenty-one (which may change if the mother changes states), did your state of residence at age fourteen have no-fault property division and alimony laws, and years of exposure to no-fault property and alimony laws since age twenty-one.[11] We are forced to restrict annual exposure measures to after the age of twenty-one because the NLSY only reports state of residence for age fourteen and then in each year from 1979-98. As a result we do not know the state of residence for women aged fifteen to twenty-two in the years before 1979.

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level of child support payments have been relatively constant since the late 1970s. This being said, we do control for family income, which includes alimony and child support payments.

[10] All dates are from Ellman and Lohr (1998).

3. Estimation and Results

Table 3 reports the percentage of the total sample engaging in 'deviant' behavior before the age of fifteen for those whose biological father is present throughout their life (180 months) and for those whose biological father is not. As is usually found, the raw numbers are quite disturbing. 27 percent of youth from non-traditional families, where the biological father is present for less than 180 months, become sexually active before age fifteen compared to only 12 percent of youth from traditional families. Similarly, youth from non-traditional families are more likely to smoke regularly, use drugs regularly, and be convicted of a crime before the age of fifteen. In contrast, there appear to be no systematic differences in alcohol consumption across family type, at least as measured by drinking once per month or more.

Raw number, however, can be quite misleading. It is possible that youth from traditional and non-traditional families are observationally different in ways beyond family structure. To properly account for this possibility, the remainder of this section examines a youth's decision to participate in deviant behavior by estimating a set of probit models that control for family background and socioeconomic factors.

3.1 **Single-Equation Probit Models**

In this section, we treat length of time exposed to one's biological father as exogenous in single-equation probit models. Therefore, these results parallel the existing literature on the effect of family structure on youth outcomes. Let the indicator variable $Y_i = 1$ if the youth participates in a specified deviant behavior before age fifteen and let $Y_i = 0$ otherwise. The choice problem is then described by the following latent variable model:

[11] We assume that women aged twenty-two in 1979 lived in the same state in 1978 and 1979. We are forced to do this because we do not know the state of residence when these women were twenty-one.

$$Y_i^* = X_i\beta_1 + D_i\delta + \varepsilon_{1i} \quad (1)$$

where Y_i^* is the net utility that a youth receives from the deviant behavior, X_i is a vector of individual characteristics (number of siblings, birth order, and gender), family characteristics (family income, mother's education, mother's race, and mother's employment status), and regional characteristics (metropolitan status and the youth's census division of residence at age fifteen), D_i is the number of months that the biological father is present in the household, and ε_{1i} is a normally distributed disturbance term with mean zero and unit variance. A youth will only participate in the deviant behavior if the expected net utility from doing so is positive, and thus the probability that the youth is observed engaging in the specified deviant behavior is given by:

$$\text{prob}(Y_i = 1) = \text{prob}(X_i\beta_1 + D_i\delta + \varepsilon_{1i} > 0) = \Phi(X_i\beta_1 + D_i\delta) \quad (2)$$

and the log likelihood function is thus given by:

$$\text{L}_1 = Y_i \ln\Phi(X_i\beta_1 + D_i\delta) + (1 - Y_i)\ln(1 - \Phi(X_i\beta_1 + D_i\delta)) \quad (3)$$

where $\Phi()$ is the standard normal cumulative distribution function and ln is the natural logarithm.

Table 4 reports the estimated determinants of smoking, drinking, sex, drug use, and conviction models. In order to more easily describe the quantitative importance of the explanatory variables, all tables report the marginal effects (∂prob(Y_i=1)/∂X_i) for continuos variables and average treatment effects for the discrete variables, in both cases evaluated at means, as well as standard errors calculated using the "delta" method. The results are largely consistent with previous findings. Youth with more educated mothers are less likely to have sex, smoke, and be convicted of a crime before the age of fifteen. For example, the child of a mother with an undergraduate degree is 8 percentage points less likely to have sex before age fifteen than an otherwise similar youth with a high school graduate mother. Male youth are 5 percentage points more likely to use drugs than female youth. Black youth are 20 and 12 percentage points less likely to smoke regularly and use marijuana regularly, respectively, than White youth. And, eldest children are substantially less likely to engage in all forms of deviant behavior compared to later born siblings. While there are few deviant behavior differences across the census divisions, youth residing in urban areas are approximately 5, 15, and 2 percentage points more likely to have sex, use marijuana regularly, and be convicted of a crime before the age of fifteen.

Also consistent with previous research, we find that the greater the percentage of a youth's life that he/she lives with their biological father, the lower the probability of smoking regularly, having sex, and being convicted of a crime. For each of these three behaviors, one extra year with the biological father decreases the probability of participation by 1 percentage point or less while 5 extra years with the biological father decreases the probability of participation in smoking, sex, and conviction by 2.5 percentage points, 4.7 percentage points, and 0.6 percentage points, respectively. In addition, we find no evidence of a correlation between the length of time that the biological father remains in the household and regular drinking or drug use on the part of teenagers.

3.2 **The Endogeneity of Family Structure**

The single-equation models presented in section 3.1 treat the length of time that the biological father remains in the household as exogenous. However, this seems unlikely to be the case. It seems much more likely that unobserved personality traits affect both parenting skills and marital status. In this case, a single-equation model will confound exposure to the biological father with unobserved parental attributes rendering biased estimates.

In this section we outline a limited dependent variable model that allows us to account for the endogeneity of time with the biological father. The choice problem is described by a two-equation model, equation (1) from section 3.1 (which is replicated here for illustrative purposes)

$$Y_i^* = X_i\beta_1 + D_i\delta + \varepsilon_{1i}$$

and,

$$D_i = X_i\beta_2 + Z_i\gamma + \varepsilon_{2i} \tag{4}$$

where Z_i is a vector of instrument(s) (a more detailed description of the instruments and their validity is provided in the discussion below), ε_{2i} is a random error term, and all other variables are as defined in section 3.1.

This limited dependent variable model with endogenous variables can be estimated using the two-stage conditional maximum likelihood (2SCML) approach proposed by Rivers and Vuong (1988).[12] In the first stage we estimate equation (4) using ordinary least squares (OLS) to obtain the least squares residual, $\hat{\varepsilon}_{2i} = D_i - \hat{\gamma}_i Z_i - \hat{\beta}_2 X_i$. In the second stage, the choice problem is then described by the following latent variable model:

$$Y_i^* = X_i\beta_1 + D_i\delta + \hat{\varepsilon}_{2i}\lambda + \varepsilon_{1i}. \tag{5}$$

A youth will only participate in the deviant behavior if the expected net utility from doing so is positive, and thus the probability that a given youth is observed engaging in the specified deviant behavior is given by:

$$\text{prob}\,(Y_i = 1) = \text{prob}\,(X_i\beta_1 + D_i\delta + \hat{\varepsilon}_{2i}\,\lambda + \varepsilon_{1i} > 0) = \Phi(X_i\beta_1 + D_i\delta + \hat{\varepsilon}_{2i}\,\lambda) \tag{6}$$

and the conditional log likelihood function is thus given by:

[12] Smith and Blundell (1986) propose a similar approach for simultaneous Tobit Models.

$$L_0 = Y_i \ln \Phi(X_i\beta_1 + D_i\delta + \hat{\varepsilon}_{2i}\,\lambda) + (1 - Y_i)\ln(1 - \Phi(X_i\beta_1 + D_i\delta + \hat{\varepsilon}_{2i}\,\lambda)). \quad (7)$$

Alternative limited information estimators, such as limited information maximum likelihood (LIML), instrumental variables probit (IVP), and generalized two-stage simultaneous probit (G2SP), could of course, be used to estimate this model.[13] While all of these estimators are strongly consistent, Rivers and Vuong (1988) note that the LIML estimator would be preferred to IVP, G2SP and 2SCML as it is the only that estimator that, in general, attains the Cramer-Rao bound. They further note that the LIML estimator is computationally burdensome and has therefore led many researchers to seek alternative estimation procedures. Rivers and Vuong (1988) argue that among the computationally less burdensome estimators (2SCML, IVP and G2SP) the 2SCML approach may be more desirable for a number of reasons. First, the 2SCML is asymptotically more efficient in some cases.[14] For example, when the model is just identified the 2SCML estimator (which in this case coincides with the LIML estimator) will attain the Cramer-Rao bound while G2SP estimator (which in this case equals the IVP estimator) generally does not. Secondly, Monte Carlo simulations show that IVP and G2SP perform less favorably than the 2SCML in small samples for both just and over identified models. Finally, the 2SCML approach (which provides an estimate of λ) allows for the construction of exogeneity test statistics that have the same asymptotic properties as the classical LIML test statistics (a detailed discussion of these tests is provided below).

In order for the model to be identified, Z_i must contain at least one variable that is not contained in X_i. An instrument is valid if it is correlated with the number of months the respondent spends with their biological father, but uncorrelated with the probability of

[13] For a more detailed discussion of these estimators see Heckman (1978), Amemiya (1978, 1979), and Lee (1981).

experiencing the youth outcome, that is, it must be uncorrelated with the error term ε_{1i}. The just identified model includes a single measure of divorce law liberalization for each mother: years of exposure to unilateral divorce laws since age twenty-one (which may change if the mother changes states). The over identified model adds three additional measures of divorce law liberalization for each mother: did your state of residence at age fourteen have a unilateral divorce law, did your state of residence at age fourteen have no-fault property division and alimony laws, and years of exposure to no-fault property division and alimony laws since age twenty-one.

Prior to 1930 every state law provided for dissolution of marriage only when there was malfeasance by the other spouse (Weitzman, 1985).[15] This legal environment began to reverse in the late 1960s when states began allowing the 'irretrievable breakdown of the marriage' to constitute grounds for divorce. Most of these reforms, broadly labeled no-fault divorce, do not require mutual consent; either party to a marriage can dissolve it at any time. Many states have also moved towards no-fault based property division and alimony rules. Using a panel of divorce rates from 1968 to 1988, Friedberg (1998) shows that liberalized divorce laws lead to rising divorce rates. Over this period the divorce rate rose from three per thousand people to five per thousand people.[16] Gruber (2000) similarly finds that individuals exposed to unilateral divorce laws are both more likely to marry early as well as get divorced. All of these findings, in addition to the results we will present later in the paper, support the view that divorce laws are correlated with the length of time that the biological father is in the household.

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[14] Newey (1987) provides a method that can be used to construct an Amemiya Generalized Least Squares (AGLS) estimator that is at least as asymptotically efficient as any two-stage estimator.

[15] Examples include abandonment, adultery, extreme cruelty, impotency, and conviction of a felony.

[16] Wolfers (2000) argues that this change results from pent up demand for divorce and that its direct effect on the divorce rate has largely run its course within eight years of a state law change.

Gruber (2000), however, argues that changes in unilateral divorce laws are not valid instruments for divorce for the following reasons: unilateral divorce laws can (1) lead to a decrease in the incidence of separation, (2) increase the incidence of marriage, and (3) change the nature of the bargaining relationship between husbands and wives. We argue that while these concerns may be particularly problematic when Census data is employed, they are less problematic when data such as the NLSY are utilized. In particular, the rich nature of the NLSY data sets allows us to control for a large set of variables, such as the employment status of the youth's mother, the education level of the youth's mother, the birth order of the youth, the number of siblings the youth has, and family income, which ensures that the remaining variation in the divorce law liberalization generates (significant) variation in the length of time exposed to the biological father (see equation (4)). For example, changes in the nature of the bargaining relationship between husbands and wives may work through the employment decisions of the wife, that is, even if the wife has a comparative advantage in housework and the husband has a comparative advantage in market work, she may choose to enter the labor market to insure against the possibility of divorce if she feels she there has been a shift in the bargaining power in the relationship.

When considering the validity of our instruments, it is also important to remember that we are not instrumenting for divorce. Rather, we are instrumenting for the length of time that the biological father is present in the household. In contrast to previous studies, our approach allows for the possibility that divorce law changes may effect not only the decision to end a marriage but also the decision to begin a marriage.

Empirically demonstrating that the instruments are correlated with number of months the biological father is present is straightforward. In the first stage regression of the determinants of the length of time that the biological father is present (see Appendix Tables A2 and A3 for the just identified and the over identified models) mother's years of exposure to unilateral divorce laws since age 21 is always individually significant at the one percent level irrespective of the youth behavior considered. The remaining instruments, with the exception of whether the mother's state of residence at age 14 had a unilateral divorce law, which is individually significant in the smoking, sex, and conviction models, are never individually significant. In the smoking, sex, and conviction models, however, the instruments are jointly significant at the one percent level, while in the drinking and marijuana models the instruments are jointly significant at the five percent level. These results tell us that after controlling for such things as the employment status of the youth's mother, the education level of the youth's mother, the birth order of the youth, the number of siblings the youth has, and family income, the remaining variation in divorce law liberalization generates significant variation in the length of time exposed to the biological father.

The next question is: are these instruments uncorrelated with youth outcomes? To explore this we re-estimate equation (3) adding the instruments to the youth outcome models. Though not a formal test, this does provide an indication of the patterns in the underlying data (see Evans and Schwab, 1995). In all cases, the proposed instruments are individually and jointly insignificant at the 10 percent level.

Of further concern is finite sample bias. Bound, Jaeger, and Baker (1993) illustrate that in finite samples, instrumental variable (IV) estimates are biased in the same direction as the ordinary least squares (OLS) estimates in over identified models. They show that the size of the bias moves in the direction of the OLS estimates as the partial R-squared (the population R-squared between the instruments and the potentially endogenous variable once the common exogenous variables have been partialled out from both the potentially endogenous variable and the instruments) moves towards zero. Furthermore they argue that if the F-statistic for the instruments from the first stage is close to unity, then finite sample bias is a concern. Although the Bound et al. (1993) methodology is not directly applicable to 2SCML, it appears that finite sample bias may be a concern in our over identified models. In particular, the first stage results have F-statistics between 2.6 and 4.3 and partial R-squareds of approximately 0.015 (see appendix Table A3). We are not, however, particularly concerned about the possibility of finite sample bias in the over identified models because the 2SCML estimates from the just identified models (which in this case are consistent and efficient as they coincide with the LIML estimators),[17] are similar in all cases (See Panels A and B of Table 5).[18] Henceforth, the discussion therefore focuses on the just identified model, although the over identified results are included for completeness.

The final concern is whether in fact the simultaneous equation approach is the desired approach. In order to ascertain this, we test the null hypothesis that the number of months that the biological father is present in the household (D_i) is exogenous (H_0: $\lambda=0$). We employ the following test statistic proposed by Rivers and Vuong (1988):

$$CLR=2[L_0-L_1] \quad (8)$$

[17] In order to determine the reliability of instrumental variable estimators, a direct comparison to LIML estimators is desired because in general, LIML estimators are consistent and efficient. For instance, Staiger and Stock (1994) find when the F-statistic is small and the number of instruments is large that LIML estimators are more reliable than standard IV estimators while Angrist, Imbens, and Krueger (1995), who develop jackknife instrumental variables estimators, find these estimators have similar finite sample properties to LIML estimators when many instruments are included.

[18] The one exception is in the conviction model. Although the magnitude of the coefficient on the length of time exposed to the biological father is similar in the just identified and over identified models, it becomes significant in the over identified model (see Panels A and B of Table 5). This may not be all that surprising given the results presented in Bound, Jaeger, and Baker (1993) illustrate that as the number of instruments increases the standard errors on the IV results are reduced while the F-statistics in the first stage deteriorate.

where CLR stands for conditional likelihood ratio, L_0 is the likelihood ratio from the unconstrained model (equation (7)) and L_1 is the likelihood ratio from the constrained model (equation (3)). The CLR statistic is distributed chi-squared with m degrees of freedom, where m is the number of endogenous variables included in equation (1).[19]

The CLR test statistics for the five youth outcomes (smoking, drinking, sex, drugs, and conviction) before age fifteen in the just identified model are reported at the bottom of Panel A of Table 5. We cannot reject the null hypothesis (H_0: λ=0) at the 10 percent significance level in the drinking and drug models, however, length of time exposed to the biological father was insignificant in the single-stage equation for these models.[20] Further, we cannot reject the null hypothesis for the conviction model, this is likely the result of noise caused by the very small fraction (4 percent) of the sample reporting a conviction before age fifteen. Finally, we reject the null hypothesis at the 10 percent significance level or less for the smoking and sex models. Therefore, the usual probit estimator of equation (1) appears to be inappropriate for the smoking and sex models and the 2SMCL approach, which takes into account the endogeneity of the number of months that the biological father was present in the household, is preferred.

Panel A of Table 5 reports the 2SCML marginal effect estimates for the just identified smoking, drinking, sex, drug use, and conviction models. Except for the coefficient on the presence of the biological father, the 2SCML results are, in general, similar to the single-equation estimates. Male youth are 5 percentage points more likely than female youth to use drugs regularly, Black youth are 27 percentage points less likely to smoke regularly than White

[19] Rivers and Vuong (1988) propose various alternative tests for exogeneity, the modified Wald statistic (MW), the Conditional Score statistic (CS), and three different Hausman statistics (M_1, M_2, and M_3). They argue that under the null hypothesis, the MW statistic, the CS statistic, the M_3 statistic, and the CLR statistic, are asymptotically equivalent.

[20] The 90 percent critical values for the chi-squared distribution is 2.706 (d.f.=1).

youth, and first born children are substantially less likely to partake in all forms of deviant behaviors (except drinking) compared to later born siblings. The main exceptions are that maternal education no longer reduces teenage promiscuity, smoking, and conviction, and Black youth are no longer less likely to use marijuana regularly than White youth.

The biggest difference between the single- and two-equation models lies in the estimated impact of time with the biological father. The magnitude of the biological father coefficient increases substantially, however, it is no longer statistically significant at the 10 percent level for conviction. The imprecision of the time with the biological father coefficient in the conviction model is again not surprising given the small number of youth reporting a conviction before age fifteen. The larger coefficients are likely the result of omitted variable bias in the OLS estimates. Omitted maternal ability is a likely candidate. As divorce laws ease, on the margin, a larger number of less able women may opt for divorce. At the same time, these lower ability women may be less able to compensate for the absence of the father.

In the single-equation model, one (five) more year(s) with the biological father decreases the probability of engaging in sexual intercourse before the age of fifteen by approximately 1.0 (4.7) percentage points. Once the endogeneity of family structure is accounted for, the estimated causal impact is a 3.9 (13.4) percentage point decrease. Similarly, one (five) more year(s) with the biological father results in a 3.8 (12.1) percentage point decrease in the probability of smoking regularly, and a 0.5 (1.6) percentage point decrease in the probability of conviction.

3.3 **Variable Definitions and Sample Selection**

To ensure that the results do not over-state the impact of exposure to the biological father due to omitted regional variables that may be correlated with youth behavior measures, Panel A in Table 6 replicates Table 5 including additional controls for current census division of residence.[21] The results presented in Panel A of Table 6 are very similar to those presented in Panel A of Table 5. For example, an additional year with the biological father reduces the probability of smoking, sex, and conviction by 4.1, 4.5, and 0.4 percentage points, respectively.

One might also be concerned that the results presented in Section 3.2 may suffer from sample selection bias because children born to young mothers disproportionately form the sample. This over-sampling occurs because the youth in the NLSY-YAS must reach age fifteen by 1998 to be included in our sample and must also have been born to women who were between the ages of fourteen and twenty-two in 1979. To check, at least in a rough way, that this is not driving the results, youth born to women who had their first child before the age of seventeen are dropped from the sample. These estimates are reported in Panel B of Table 6. Again the results are quite similar. An additional five years with the biological father reduces the probability of becoming sexually active by 12.9 percentage points, reduces the probability of smoking regularly by 12.3 percentage points, and reduces the probability of being convicted by 1.5 percentage points. Interestingly, the coefficient on length of time exposed to biological father is now statistically significant at the 10 percent level or less in the conviction model. This illustrates the large degree of noise associated with the conviction model due to the small fraction of youth's who were convicted of a crime.

Number of siblings may also be endogenous if some couples try to 'save' their marriage by having an additional child. To ensure that our results are robust to this possibility, Panel C reports the coefficient for time with the biological father for models that exclude number of siblings from the list of explanatory variables. Once again the results are similar.

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[21] Ideally, we would control for mother's state of residence at age fourteen and the current state of residence, however, the cell sizes become too small in many cases.

One may conversely believe that the results are driven by the omission of parental behavior measures from the explanatory variable list. While we do not have access to adequate measures of sexual behavior and conviction for mothers and fathers, the NLSY does report smoking behavior for mothers in 1998. As can be seen from Panel D, the results are in fact robust to the inclusion of this measure.[22]

Finally, one might also be concerned that the results are sensitive to the presence of a father definition. As discussed in section 2, a more restrictive definition of the presence of a father could be used, similarly a more expansive definition could also be employed. The first four columns of Panel A of Table 7 present 4 different father definitions for the just identified model. In columns one through four a man is considered to be the youth's father if he was married to the mother within 0, 12, 36, and 60 months of the youth's birth, respectively. If the father marries the mother at or before the youth's birth, the number of months that the father is present is measured by the number of months that the marriage lasts. If, on the other hand, the father marries the mother after the youth's birth, the number of months that the father is present is measured by the number of months that the marriage lasts plus the number of months from the youth's birth until the marriage began. In all other cases, we assume that the youth never lives with his/her biological father. Column 3 of Table 7 represents the father definition used in the previous sections of the paper. Regardless of the father definition utilized, the results are again very similar. For instance, depending on the definition utilized, an additional year with the biological father reduces the probability of becoming sexually active by 3.6 percentage points to

[22] The results are also similar when controls for the frequency of the father's visits are included and if the small number of cases where the deviant behavior takes place before the father leaves the household are excluded. These results are available from the authors upon request.

4.2 percentage points, with the definition of biological father that we chose lying in the middle at 3.7 percentage points.

An additional concern with the presence of father definition is that the crucial issue may be the presence of a father-figure rather than a biological relationship. In other words, it is possible that a stepfather has the same deterrence effect for deviant behavior as a biological father. This may be particularly important for some youths, who by definition are defined as not having a biological father, that is, youths whose mother's married after 36 months of their birth. As described in section 2, this group of youths on average has their first stepfather by age 7 and lives with this first stepfather for 77 months. Therefore, excluding time spent with stepfathers may bias the estimates downward as we may be attributing good outcomes to youth defined as not having a "father".

Columns 5 through 8 of Panel A in Table 7 replaces months that the mother is married to the biological father (using the four definition defined above) with total months when either a biological father or a stepfather is present. As the results for all definitions are similar, we focus on column seven, which builds on the father definition used throughout the paper. The estimated impact of the presence of a "father" are somewhat larger for all models. One more year with some type of father results in 5.1, 4.9, and 0.6 percentage point decreases in the probability of smoking regularly, having sex, and being convicted of a crime before the age of fifteen, respectively. Thus, excluding time spent with stepfathers does appear to bias the estimates downward.

4. Conclusion

Prior to the enactment of no-fault divorce laws, a marriage could not be legally dissolved without cause even in the event that both parties wanted it to end. Today, in contrast, most states have no-fault divorce laws that make it impossible for one party to keep a marriage intact if the other party wishes to dissolve it. Much of the criticism of the move towards unilateral divorce is that children's interests are not protected. This argument, as well as attempts to return to fault-based laws and convent marriage reforms in some states, clearly rest on the presumption that children are harmed by exposure to single-parent arrangements. Despite the myriad of studies providing evidence of the correlation between divorce and poor youth outcomes, to date there is limited evidence to support a claim of causation.

In contrast to most previous research, this study provides substantial evidence of a causal relationship between family dissolution and poor youth outcomes. Stated somewhat differently, we have shown that children are in fact harmed by reduced time with their biological father. Children who spend part of their childhood/youth living in a household that does not include their biological father are more likely to smoke regularly, become sexually active, and be convicted of a crime.

We believe that the results presented in this study should be viewed as a first step towards understanding the impact of family structure on youth outcomes. While children residing with mothers constitute the majority of children in the United States today, it is nonetheless important that future work expands the sample to include children raised by fathers and other types of alternate caregivers. However, this endeavor requires more extensive data than are presently available.

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Table 1. Summary Statistics

	Sample Size	Mean	Standard Deviation
Smoking	1258	0.1806	0.3849
Drinking	803	0.1502	0.3575
Sex	1297	0.1927	0.3945
Marijuana Use	861	0.1596	0.3665
Conviction	1282	0.0397	0.1953
Months with Biological Father	1297	116.5925	76.8073
Number of Siblings	1297	2.6459	1.1376
First Born Child	1297	0.6323	0.4824
Male	1297	0.5051	0.5002
Black	1297	0.2019	0.4016
Race other than White or Black	1297	0.0266	0.1610
Mother's Years of Education	1297	12.2463	1.7376
Mother's Average Hours of Work	1297	24.0838	15.9283
Average Net Family Income	1297	45.1704	49.7098
SMSA, Not Central City	1297	0.3890	0.4877
SMSA, Central City Not Known	1297	0.2128	0.4095
SMSA, Central City	1297	0.1123	0.3159

All youth outcomes measure participation before age fifteen. Means and standard deviations calculated using 1998 youth sampling weights.

Table 2a. Timing and Duration of the Mother's First Marriage Experienced by Youth

	% of Sample	Mean Age at Marriage (in Months)	Mean Age at Divorce (in Months)
Mother's Marriage to Biological Father and/or First Husband After the Youth's Birth			
Married at Youth's Birth	0.73	0	145
Not Married at Youth's Birth	0.27	NA	NA
Married During Youth's Life	0.17	55	153
Married at or Before Youth is 36 Months	0.07	16	140
Married After Youth is 36 Months	0.10	85	162
Never Marries	0.10	NA	NA

Table 2b. Timing and Duration of Fathers

	% of Sample	Mean Age at Father's Arrival (in Months)	Mean Months with Father (in Months)	% of Group with Stepfather	Mean Age at Stepfather's Arrival (in Months)	Mean Months with Stepfather (in Months)
Time with Father and Stepfather						
Married at Youth's Birth	0.73	0	145	0.20	99	67
Youth 0 Months < Marriage <= Youth 36 Months	0.07	16	140	0.26	117	55
Marriage > Youth 36 Months	0.10	NA	NA	NA	85	77

Table 3. Participation in Deviant Behavior

	<100% of Life with Biological Father	100% of Life with Biological Father
Smoking	**0.1984**	0.1671
Drinking	0.1523	0.1521
Sex	**0.2708**	0.1239
Marijuana	**0.1790**	0.1393
Conviction	**0.0508**	0.0299

All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used. Bold indicates probabilities greater than those for traditional families (100% of life with father) at the 10% level or better.

Table 4. Single-Equation Probit Estimates

	Smoking	Drinking	Sex	Drugs	Conviction
Months with Biological Father	**-0.0005**	0.0001	**-0.0009**	-0.0003	**-0.0001**
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0001)
Number of Siblings	-0.0169	0.0072	**-0.0199**	-0.0119	**0.0056**
	(0.0126)	(0.0148)	(0.0119)	(0.0138)	(0.0035)
First Born	**-0.1085**	**-0.0779**	**-0.0675**	**-0.0797**	**-0.0245**
	(0.0315)	(0.0346)	(0.0287)	(0.0357)	(0.0108)
Male	0.0069	0.0113	0.0041	**0.0514**	0.0106
	(0.0260)	(0.0301)	(0.0253)	(0.0290)	(0.0083)
Black	**-0.1955**	-0.0468	-0.0182	**-0.1186**	-0.0119
	(0.0195)	(0.0339)	(0.0309)	(0.0258)	(0.0073)
Race other than White or Black	-0.0389	**0.1448**	0.0123	0.0518	**0.0617**
	(0.0589)	(0.1018)	(0.0681)	(0.0870)	(0.0477)
Mother's Years of Education	**-0.0227**	-0.0027	**-0.0199**	-0.0091	**-0.0054**
	(0.0080)	(0.0075)	(0.0069)	(0.0088)	(0.0021)
Mother's Average Hours of Work	-0.0009	**0.0016**	-0.0001	0.0009	0.0000
	(0.0009)	(0.0010)	(0.0008)	(0.0010)	(0.0002)
Average Net Family Income	0.0002	0.0001	-0.0001	0.0001	-0.0002
	(0.0003)	(0.0003)	(0.0003)	(0.0003)	(0.0001)
Impact of One more Year with Father	-0.5%	0.1%	-1.0%	-0.4%	-0.1%
Impact of Five more Years with Father	-2.5%	0.4%	-4.7%	-1.9%	-0.6%
Sample Size	1258	803	1297	861	1282

All models also include the youth's census division of residence at age fifteen indicator variables and metropolitan status indicator variables. Heteroskedastic consistent standard errors are in parentheses. Bold coefficients are statistically significant at the 10% level. All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used.

Table 5. Simultaneous Equation Probit Estimates

	Smoking	Drinking	Sex	Marijuana	Conviction
Panel A: Just Identified					
Months with Biological Father	**-0.0036**	0.0012	**-0.0035**	-0.0007	-0.0005
	(0.0019)	(0.0019)	(0.0019)	(0.0017)	(0.0005)
Number of Siblings	-0.0024	0.0039	-0.0113	-0.0112	**0.0070**
	(0.0148)	(0.0151)	(0.0133)	(0.0144)	(0.0038)
First Born	**-0.1491**	-0.0648	**-0.1016**	**-0.0861**	**-0.0294**
	(0.0425)	(0.0420)	(0.0374)	(0.0480)	(0.0137)
Male	0.0276	0.0006	0.0214	**0.0530**	0.0129
	(0.0279)	(0.0358)	(0.0273)	(0.0300)	(0.0082)
Black	**-0.2726**	0.0340	-0.1629	-0.1331	-0.0265
	(0.0480)	(0.1558)	(0.0813)	(0.0727)	(0.0179)
Race other than White or Black	-0.0782	0.2130	-0.0560	0.0381	0.0357
	(0.0477)	(0.1612)	(0.0665)	(0.1075)	(0.0478)
Mother's Years of Education	-0.0069	-0.0084	-0.0058	-0.0077	-0.0035
	(0.0129)	(0.0129)	(0.0118)	(0.0117)	(0.0030)
Mother's Average Hours of Work	-0.0013	**0.0017**	-0.0005	0.0008	0.0000
	(0.0009)	(0.0010)	(0.0009)	(0.0010)	(0.0003)
Average Net Family Income	0.0006	0.0000	0.0003	0.0001	-0.0002
	(0.0004)	(0.0004)	(0.0004)	(0.0003)	(0.0001)
CLR	4.4675	0.5767	2.8879	0.0643	0.5280
Impact of One more Year with Father	-0.0384	0.0151	-0.0386	-0.0080	-0.0050
Impact of Five more Years with Father	-0.1209	0.0862	-0.1335	-0.0365	-0.0158
Panel B: Over Identified					
Months with Biological Father	**-0.0026**	0.0012	**-0.0031**	-0.0003	**-0.0008**
	(0.0014)	(0.0016)	(0.0015)	(0.0014)	(0.0004)
Number of Siblings	-0.0068	0.0039	-0.0128	-0.0119	**0.0080**
	(0.0135)	(0.0149)	(0.0129)	(0.0141)	(0.0037)
First Born	**-0.1362**	**-0.0652**	**-0.0968**	**-0.0796**	**-0.0336**
	(0.0384)	(0.0404)	(0.0343)	(0.0443)	(0.0131)
Male	0.0216	0.0010	0.0180	**0.0514**	**0.0144**
	(0.0273)	(0.0344)	(0.0268)	(0.0297)	(0.0085)
Black	**-0.2501**	0.0304	-0.1454	-0.1183	**-0.0350**
	(0.0377)	(0.1345)	(0.0714)	(0.0661)	(0.0143)
Race other than White or Black	-0.0667	**0.2099**	-0.0466	0.0521	0.0195
	(0.0514)	(0.1514)	(0.0643)	(0.1072)	(0.0365)
Mother's Years of Education	-0.0119	-0.0081	-0.0080	-0.0091	-0.0017
	(0.0110)	(0.0118)	(0.0106)	(0.0109)	(0.0026)
Mother's Average Hours of Work	-0.0012	**0.0017**	-0.0004	0.0009	-0.0001
	(0.0009)	(0.0010)	(0.0009)	(0.0010)	(0.0002)
Average Net Family Income	0.0005	0.0000	0.0002	0.0001	-0.0001
	(0.0004)	(0.0003)	(0.0004)	(0.0003)	(0.0001)
CLR	4.1702	0.7238	3.3741	0.0023	2.8251
Impact of One more Year with Father	-0.0288	0.0145	-0.0346	-0.0038	-0.0076
Impact of Five more Years with Father	-0.1034	0.0823	-0.1249	-0.0182	-0.0182
Sample Size	1258	803	1297	861	1282

All models also include the youth's census division of residence at age fifteen indicator variables and metropolitan status indicator variables. Heteroskedastic consistent standard errors are in parentheses. Bold coefficients are statistically significant at the 10% level. All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used.

Table 6. Simultaneous Equation Probit Estimates (Alternative Specifications)

	Smoking	Drinking	Sex	Marijuana	Conviction
Panel A - Adding Census Division of Residence when Youth is age 15					
Just Identified Model					
Months with Biological Father	**-0.0038**	0.0011	**-0.0042**	-0.0012	-0.0004
	(0.0019)	(0.0019)	(0.0019)	(0.0018)	(0.0004)
Impact of One more Year with Father	-0.0410	0.0139	-0.0454	-0.0143	-0.0040
Impact of Five more Years with Father	-0.1217	0.0793	-0.1435	-0.0602	-0.0128
Over Identified Model					
Months with Biological Father	**-0.0037**	0.0011	**-0.0034**	-0.0004	-0.0004
	(0.0021)	(0.0020)	(0.0020)	(0.0019)	(0.0005)
Impact of One more Year with Father	-0.0327	0.0044	-0.0347	-0.0112	-0.0064
Impact of Five more Years with Father	-0.1094	0.0228	-0.1243	-0.0491	-0.0151
Sample Size	1258	803	1297	861	1282
Panel B - Excluding Mother's Whose First Birth is Before Age 17					
Just Identified Model					
Months with Biological Father	**-0.0030**	0.0004	**-0.0031**	-0.0010	**-0.0007**
	(0.0015)	(0.0016)	(0.0016)	(0.0016)	(0.0004)
Impact of One more Year with Father	-0.0402	-0.0141	-0.0378	-0.0048	-0.0047
Impact of Five more Years with Father	-0.1230	0.0796	-0.1294	-0.0228	-0.0153
Over Identified Model					
Months with Biological Father	**-0.0030**	0.0009	**-0.0032**	-0.0004	**-0.0007**
	(0.0016)	(0.0018)	(0.0016)	(0.0016)	(0.0004)
Impact of One more Year with Father	-0.0331	0.0117	-0.0353	-0.0052	-0.0071
Impact of Five more Years with Father	-0.1118	0.0646	-0.1243	-0.0247	-0.0179
Sample Size	1183	743	1218	785	1203

All models also include indicator variables for mother's census division of residence at age fourteen, the youth's census division of residence at age fifteen, and metropolitan status indicator variables. Heteroskedastic consistent standard errors are in parentheses. Bold coefficients are statistically significant at the 10% level. All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used.

Table 6. Simultaneous Equation Probit Estimates (Alternative Specifications) - Continued

	Smoking	Drinking	Sex	Marijuana	Conviction
Panel C - Excluding Number of Siblings Variable					
Just Identified Model					
Months with Biological Father	**-0.0036**	0.0012	**-0.0035**	-0.0007	-0.0005
	(0.0019)	(0.0019)	(0.0019)	(0.0017)	(0.0005)
Impact of One more Year with Father	-0.0388	0.0153	-0.0387	-0.0078	-0.0047
Impact of Five more Years with Father	-0.1216	0.0876	-0.1341	-0.0359	-0.0365
Over Identified Model					
Months with Biological Father	**-0.0026**	0.0012	**-0.0030**	-0.0003	**-0.0008**
	(0.0014)	(0.0016)	(0.0015)	(0.0014)	(0.0004)
Impact of One more Year with Father	-0.0292	0.0146	-0.0341	-0.0037	-0.0079
Impact of Five more Years with Father	-0.1043	0.0830	-0.1242	-0.0180	-0.0189
Sample Size	1258	803	1297	861	1282
Panel D - Including Mother's Smoking Status in 1998					
Just Identified Model					
Months with Biological Father	**-0.0034**	0.0015	**-0.0034**	-0.0006	-0.0004
	(0.0020)	(0.0020)	(0.0019)	(0.0018)	(0.0005)
Impact of One more Year with Father	-0.0371	0.0185	-0.0373	-0.0720	-0.0045
Impact of Five more Years with Father	-0.1180	0.1085	-0.1298	-0.0331	-0.0143
Over Identified Model					
Months with Biological Father	**-0.0024**	0.0013	**-0.0029**	-0.0003	**-0.0007**
	(0.0014)	(0.0017)	(0.0015)	(0.0015)	(0.0004)
Impact of One more Year with Father	-0.0271	0.0156	-0.0328	-0.0031	-0.0069
Impact of Five more Years with Father	-0.0990	0.0896	-0.1202	-0.0152	-0.0167
Sample Size	1258	803	1297	861	1282

All models also include indicator variables for mother's census division of residence at age fourteen, the youth's census division of residence at age fifteen, and metropolitan status indicator variables. Heteroskedastic consistent standard errors are in parentheses. Bold coefficients are statistically significant at the 10% level. All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used.

Table 7. Simultaneous Equation Probit Estimates (Alternative Father Definitions)

	Father Definitions				Father plus Stepfather Definitions			
	0 Months	12 Months	36 Months	60 Months	0 Months	12 Months	36 Months	60 Months
Panel A: Just Identified Model								
Smoking	**-0.0032**	**-0.0038**	**-0.0036**	**-0.0036**	**-0.0049**	**-0.0049**	**-0.0049**	**-0.0050**
	(0.0017)	(0.0021)	(0.0019)	(0.0020)	(0.0026)	(0.0027)	(0.0026)	(0.0027)
One more Year with Father	-0.0349	-0.0406	-0.0384	-0.0384	-0.0507	-0.0510	-0.0508	-0.0513
Five more Years with Father	-0.1156	-0.1240	-0.1209	-0.1203	-0.1337	-0.1339	-0.1338	-0.1340
Drinking	0.0013	0.0014	0.0012	0.0011	0.0016	0.0016	0.0016	0.0015
	(0.0020)	(0.0022)	(0.0019)	(0.0017)	(0.0025)	(0.0025)	(0.0024)	(0.0024)
One more Year with Father	0.0158	0.0174	0.0151	0.0139	0.0199	0.0200	0.0196	0.0194
Five more Years with Father	0.0903	0.1010	0.0862	0.0785	0.1174	0.1180	0.1157	0.1143
Sex	**-0.0032**	**-0.0038**	**-0.0035**	**-0.0034**	**-0.0045**	**-0.0045**	**-0.0045**	**-0.0045**
	(0.0017)	(0.0021)	(0.0019)	(0.0018)	(0.0024)	(0.0025)	(0.0024)	(0.0025)
One more Year with Father	-0.0357	-0.0421	-0.0386	-0.0372	-0.0487	-0.0490	-0.0487	-0.0486
Five more Years with Father	-0.1269	-0.1397	-0.1335	-0.1309	-0.1508	-0.1513	-0.1509	-0.1511
Marijuana	-0.0007	-0.0008	-0.0007	-0.0006	-0.0008	-0.0008	-0.0008	-0.0008
	(0.0018)	(0.0020)	(0.0017)	(0.0017)	(0.0022)	(0.0022)	(0.0022)	(0.0022)
One more Year with Father	-0.0081	-0.0088	-0.0080	-0.0073	-0.0097	-0.0097	-0.0097	-0.0096
Five more Years with Father	-0.0370	-0.0400	-0.0365	-0.0338	-0.0433	-0.0436	-0.0436	-0.0432
Conviction	-0.0005	-0.0005	-0.0005	-0.0005	-0.0006	-0.0006	-0.0006	-0.0006
	(0.0005)	(0.0006)	(0.0005)	(0.0005)	(0.0007)	(0.0007)	(0.0007)	(0.0007)
One more Year with Father	-0.0048	-0.0054	-0.0050	-0.0049	-0.0064	-0.0065	-0.0064	-0.0065
Five more Years with Father	-0.0156	-0.0166	-0.0158	-0.0156	-0.0176	-0.0177	-0.0176	-0.0176
Panel B: Over Identified Model								
Sex	**-0.0020**	**-0.0021**	**-0.0026**	**-0.0025**	**-0.0037**	**-0.0037**	**-0.0039**	**-0.0039**
	(0.0011)	(0.0013)	(0.0014)	(0.0014)	(0.0020)	(0.0020)	(0.0020)	(0.0021)
One more Year with Father	-0.0222	-0.0240	0.0288	-0.0281	-0.0399	-0.0402	-0.0417	-0.0420
Five more Years with Father	-0.0870	-0.0918	-0.1034	-0.1014	-0.1228	-0.1232	-0.1251	-0.1253
Drinking	0.0009	0.0009	0.0012	0.0011	0.0010	0.0010	0.0011	0.0011
	(0.0014)	(0.0016)	(0.0016)	(0.0014)	(0.0021)	(0.0021)	(0.0021)	(0.0021)
One more Year with Father	0.0110	0.0108	0.0145	0.0141	0.0125	0.0125	0.0130	0.0137
Five more Years with Father	0.0609	0.0593	0.0823	0.0799	0.0698	0.0699	0.0732	0.0773
Smoking	**-0.0026**	**-0.0028**	**-0.0031**	**-0.0030**	**-0.0046**	**-0.0047**	**-0.0048**	**-0.0048**
	(0.0012)	(0.0013)	(0.0015)	(0.0014)	(0.0021)	(0.0021)	(0.0021)	(0.0021)
One more Year with Father	-0.0293	-0.0315	-0.0346	-0.0338	-0.0499	-0.0503	-0.0511	-0.0513
Five more Years with Father	-0.1117	-0.1175	-0.1249	-0.1233	-0.1519	-0.1525	-0.1536	-0.1541
Marijuana	-0.0002	-0.0004	-0.0003	-0.0002	-0.0005	-0.0005	-0.0005	-0.0004
	(0.0012)	(0.0015)	(0.0014)	(0.0013)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
One more Year with Father	-0.0024	-0.0044	-0.0038	-0.0019	-0.0059	-0.0061	-0.0056	-0.0048
Five more Years with Father	-0.0118	-0.0210	-0.0182	-0.0092	-0.0277	-0.0285	-0.0265	-0.0227
Conviction	**-0.0007**	**-0.0008**	**-0.0008**	**-0.0009**	**-0.0010**	**-0.0010**	**-0.0010**	**-0.0010**
	(0.0004)	(0.0004)	(0.0004)	(0.0004)	(0.0006)	(0.0006)	(0.0006)	(0.0006)
One more Year with Father	-0.0071	-0.0076	-0.0076	-0.0082	-0.0091	-0.0092	-0.0089	-0.0094
Five more Years with Father	-0.0178	-0.0182	-0.0182	-0.0182	-0.0188	-0.0189	-0.0188	-0.0188

All models include the variables listed in Table 5. Heteroskedastic consistent standard errors are in parentheses. Bold coefficients are statistically significant at the 10% level. All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used.

Appendix Table A1. State Divorce Law Changes

	Irretrievable Breakdown (No-Fault Divorce)	No-Fault Based Property Division and Alimony
Alabama	1971	
Alaska	1974	1974
Arizona	1973	1973
Arkansas	1979	1979
California	1969	1969
Colorado	1971	1971
Connecticut	1973	
Delaware	1974	1974
Florida	1971	1986
Georgia	1973	
Hawaii	1972	1960
Idaho	1971	1990
Illinois	1983	1977
Indiana	1973	1973
Iowa	1970	1972
Kansas		1990
Kentucky	1972	
Louisiana	1973	
Maine		1985
Maryland		
Massachusetts		
Michigan	1971	
Minnesota	1974	1974
Mississippi	1976	
Missouri	1973	
Montana	1975	1975
Nebraska	1972	1972
Nevada	1931	
New Hampshire		
New Jersey		1980
New Mexico		1976
New York		
North Carolina		
North Dakota	1971	
Ohio	1974	
Oklahoma		1975
Oregon	1971	1971
Pennsylvania	1980	
Rhode Island	1975	
South Carolina		
South Dakota	1985	
Tennessee	1977	
Texas	1969	
Utah	1987	1987
Vermont		
Virgina		
Washington	1973	1973
West Virginia	1977	
Wisconsin	1977	1977
Wyoming	1977	

These dates are all from Ellman and Lohr (1998).

Appendix Table A2. First Stage Regression (Just Identified)

	Smoking	Drinking	Sex	Marijuana	Conviction
Number of Siblings	**4.5758**	2.7755	3.2883	2.0653	3.6444
	(2.3986)	(3.2432)	(2.3722)	(3.0665)	(2.4159)
First Born	**-9.8278**	-8.8539	**-10.4592**	**-14.3443**	**-8.0182**
	(4.9519)	(6.2109)	(4.8279)	(5.9933)	(4.9206)
Male	6.4921	**9.7046**	6.6656	4.6021	6.2552
	(4.6288)	(5.8363)	(4.5470)	(5.7738)	(4.6112)
Black	**-73.2512**	**-72.7452**	**-72.2382**	**-70.6923**	**-71.8756**
	(5.2339)	(6.6379)	(5.1777)	(6.3803)	(5.2073)
Race other than White or Black	-16.9216	**-35.3446**	**-26.9473**	**-31.3946**	**-25.3226**
	(16.0259)	(18.3539)	(15.3626)	(19.0173)	(15.3145)
Mother's Years of Education	**4.8849**	**4.7830**	**5.0859**	**3.6658**	**4.6615**
	(1.4806)	(1.8069)	(1.4458)	(1.8094)	(1.4391)
Mother's Average Hours of Work	-0.1247	-0.0864	-0.1175	-0.1123	-0.1366
	(0.1540)	(0.1983)	(0.1498)	(0.1897)	(0.1514)
Average Net Family Income	**0.1450**	**0.1047**	**0.1457**	**0.1063**	**0.1452**
	(0.0669)	(0.0635)	(0.0636)	(0.0635)	(0.0629)
Instrument:					
Years of Exposure to Unilateral Divorce	**1.3094**	**1.6206**	**1.3200**	**1.7110**	**1.4150**
Laws Since Age 21	(0.4393)	(0.5603)	(0.4273)	(0.5682)	(0.4251)
F-Statistic for Instrument	**8.8800**	**8.3700**	**9.5400**	**9.0700**	**11.0800**
P-Value of F-Statistic for Instrument	0.0029	0.0039	0.0021	0.0027	0.0009
R-Squared	0.2188	0.2025	0.2218	0.2253	0.2159
Partial R-Squared	0.0076	0.0110	0.0078	0.0116	0.0090
Sample Size	1258	803	1297	861	1282

All models also include the youth's census division of residence at age fifteen indicator variables and metropolitan status indicator variables. Heteroskedastic consistent standard errors are in parentheses. Bold coefficients are statistically significant at the 10% level. All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used.

Appendix Table A3. First Stage Regression (Over Identified)

	Smoking	Drinking	Sex	Marijuana	Conviction
Number of Siblings	**4.7434**	2.9440	3.4439	2.1429	3.8085
	(2.3855)	(3.2232)	(2.3602)	(3.0451)	(2.4043)
First Born	**-8.2402**	-7.6080	**-9.1232**	**-12.8084**	-6.5637
	(4.9468)	(6.2187)	(4.8333)	(6.0190)	(4.9249)
Male	5.6720	9.2085	**5.7973**	4.1406	5.5248
	(4.6322)	(5.8424)	(4.5538)	(5.7748)	(4.6166)
Black	**-73.3355**	**-73.2477**	**-72.7089**	**-71.0861**	**-71.8633**
	(5.2610)	(6.7095)	(5.1919)	(6.4642)	(5.2222)
Race other than White or Black	-15.1486	**-33.0315**	**-25.4270**	-28.7415	-24.2463
	(16.2556)	(18.8235)	(15.5845)	(19.4563)	(15.5995)
Mother's Years of Education	**4.8601**	**4.7670**	**5.0584**	**3.6452**	**4.6093**
	(1.4839)	(1.8073)	(1.4484)	(1.8056)	(1.4418)
Mother's Average Hours of Work	-0.1367	-0.0757	-0.1236	-0.1148	-0.1436
	(0.1567)	(0.2013)	(0.1526)	(0.1939)	(0.1536)
Average Net Family Income	**0.1445**	0.1036	**0.1455**	**0.1041**	**0.1440**
	(0.0667)	(0.0635)	(0.0636)	(0.0632)	(0.0626)
Instruments:					
Years of Exposure to Unilateral Divorce Laws Since Age 21	**1.7434**	**1.7960**	**1.6392**	**1.9020**	**1.8011**
	(0.4715)	(0.6028)	(0.4616)	(0.6112)	(0.4589)
State of Residence at Age 14 Have Unilateral Divorce Laws	**-15.2341**	-7.4713	**-11.7298**	-7.5632	**-13.5645**
	(6.3115)	(7.8916)	(6.2232)	(7.7733)	(6.3273)
Years of Exposure to No Fault Property Division and Alimony Laws Since Age 21	-0.4617	-0.2253	-0.3984	-0.0459	-0.3287
	(0.4583)	(0.5989)	(0.4552)	(0.5862)	(0.4547)
State of Residence at Age 14 Have No Fault Property Division and Alimony Laws	-1.1469	-8.1228	-4.9160	-12.3487	-0.7352
	(9.6456)	(11.2461)	(9.5917)	(11.3636)	(9.6828)
F-Statistic for Instruments	**4.2300**	**2.6000**	**3.8600**	**3.1300**	**4.3100**
P-Value of F-Statistic for Instruments	0.0021	0.0348	0.0040	0.0144	0.0018
R-Squared	0.2259	0.2064	0.2272	0.2303	0.2213
Partial R-Squared	0.0147	0.0149	0.0133	0.0166	0.0143
Sample Size	1258	803	1297	861	1282

All models also include the youth's census division of residence at age fifteen indicator variables and metropolitan status indicator variables. Heteroskedastic consistent standard errors are in parentheses. Bold coefficients are statistically significant at the 10% level. All youth outcomes measure participation before age fifteen. 1998 youth sampling weights are used.