IS INTERGENERATIONAL EARNINGS MOBILITY AFFECTED BY DIVORCE?

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Abstract

This study examines whether the intergenerational transmission of human capital, measured by intergenerational earnings mobility, is affected by divorce. Using the Panel Study of Income Dynamics, I find that, with each additional year in a family involving a single or a step parent, the earnings mobility between biological fathers and children rises and the mobility between mothers and daughters falls. However, using either sibling fixed effects or instrumental variable estimation, I find that the association between family structure and father-child mobility is explained by selection. These findings have two important implications. First, they imply that the increase in father-son mobility observed in other studies can be explained by the rise in single and step parent families over the same period. Second, these findings imply that the connection between fathers and children would have been weak whether or not a divorce occurred, which does not support the hypothesis that father absence is an important factor contributing to differences in child outcomes across family structures.

JEL Codes: J62, J12

Keywords: intergenerational earnings mobility, family structure

1 Introduction

According to Bumpass and Lu (2000), nearly half of all children born in the United States today will experience a single or step parent family at some point in their childhood. These children are more likely to drop out of high school and less likely to attend college (McLanahan & Bumpass, 1988; Grogger & Ronan, 1995; Ginther & Pollak, 2000). Girls from single parent families are more likely to become teen mothers (McLanahan & Sandefur, 1994) and boys from these family structures are more likely to be idle – not working or in school – in their early twenties (Haveman & Wolfe, 1994; McLanahan, 1997). There is an active discussion in the sociological and economic literatures about the mechanisms by which family structure affects these children. One argument – the father-absence hypothesis – is that having a non-resident biological father reduces access to the father's time, reduces the amount of time spent with any parent, reduces the amount of supervision, and increases the level of stress in the child's household (Popenoe, 1996; Murray, 1995). In this paper, I look for support for the father-absence hypothesis by examining whether the intergenerational transmission of human capital – proxied by intergenerational earnings mobility – is weaker in single and step parent families.

Intergenerational earnings *immobility*, measured by the elasticity of child's earnings with respect to parent's earnings, signifies the importance of family background in determining a child's earnings.[1] That which family earnings cannot explain is called intergenerational *mobility*. If fathers have a diminished capacity to influence their children when they do not reside with them, which could contribute to differences in child outcomes across family structures, mobility will be higher in families with a non-resident father. However, because divorce is fraught with choices, mobility could also be higher in single and step parent families if fathers who are less likely to influence their children are also more likely to divorce. I address this selection issue in two ways. First, I examine

[1] This elasticity is often loosely referred to as the correlation between the parents' and children's earnings, although this is only technically true if the variance of the parents' and the children's earnings distributions are equal.

sibling differences in single and step family experience using fixed effects. Second, I instrument the number of years in a single or step parent family with the child's exposure to no-fault divorce laws. Using data on the earnings of parents and children from the Panel Study of Income Dynamics (PSID), I find that with each additional year in either a single-parent or step family, the earnings mobility between biological fathers and their children rises, but that this association is explained by selection into family structure. That is, the connection between fathers and children would have been weak whether or not a divorce occurred, which suggests that one consequence of divorce – father absence – is not responsible for differences in mobility across family structures.

The correlation between family structure and mobility is of additional interest because it could explain at least in part why intergenerational mobility between fathers and sons has increased in the US over the last few decades (Corcoran, 2001; Fertig, 2003; Harding, Jencks, Lopoo, & Mayer, 2002; Mayer & Lopoo, 2001).[2] Intergenerational mobility is often used as a measure of *equality of opportunity* because it represents the degree to which one's earnings are not pre-determined by his/her parent's earnings. Thus, the trend in father-son mobility may be interpreted as an increase in equality of opportunity. Policymakers would like to believe that education programs or redistributive policies are responsible. However, if mobility is higher in single and step parents families, and the proportion of these family types are growing, then overall mobility could be rising because of a compositional change in the make-up of American families.

The paper is organized as follows. I describe the data and the sample in the next section. In section 3, I examine the effect of family structure on *earnings* mobility without accounting for selection. I first develop a model which motivates the empirical strategy and then present the findings. In section 4, I present results using hourly wage and education to lend support to the interpretation of the earnings regressions in the previous section. In section 5, I account for

[2]In contrast to the studies just cited, Levine and Mazumder (2002) use the National Longitudinal Surveys (NLS), the GSS, and the PSID to conclude that family income mobility has significantly decreased between 1980 and the early 1990s.

selection into family structure using two techniques to argue that the association between father-child mobility and family structure is not causal. Finally, in section 6, I divide the sample into five cohorts and show that father-son mobility does not rise significantly when family structure controls are included in the specification.

2 Data and Sample

The data used in this study come from the core sample of the PSID, also called the Survey Research Center sample.[3] The PSID is well-suited for both intergenerational and sibling analysis because families can be easily matched across and within generations. Although the motivations for this paper center around fathers, I am also interested in the mobility with respect to mothers. If the investment of a parent's time is important, then the greater amount of time spent with a mother relative to a father in a single-parent family would make mother-child mobility lower. Thus, I consider both biological parents and children. I separate sons and daughters in the main analysis since family structure may have different effects by gender (Powell & Parcel, 1997) and the child's gender may affect his or her mobility. For the sibling fixed effects analysis, the sample includes parents and their children of either sex because I do not want to impose a same-sex sibling restriction on the sample. I also pool boys and girls for the instrumental variables analysis to increase sample size.

Summary statistics on the sample of parents and sons and the sample of parents and daughters are presented in Table 1. Family structure is characterized in this paper as the number of years out of 18 that a child lives with both biological parents, with a single biological parent, and/or

[3]Because this sample was constructed as an equal probability sample, I do not include weights in this analysis. There are concerns with regard to attrition given the length of the panel, however, Fitzgerald, Gottschalk, and Moffitt (1998) analyzed the impacts of sample attrition in the PSID and did not find evidence of attrition bias in the intergenerational earnings relationship. Still, one might argue that a disproportionate number of alternative families are lost to attrition since they have a high rate of residential mobility (Astone & McLanahan, 1994).

with a biological parent and a step-parent.[4] Hereafter I will often refer to the latter two family types as alternative families. Given that Wolfe, Haveman, Ginther, and An (1996) find that one-year 'window' variables can lead to unreliable estimates, the duration in an alternative family is a more appropriate approximation of family structure.[5] Because of greater variation, it also has the advantage of greater precision. For each sample, three columns provide statistics on mutually exclusive groups: those children who have never experienced an alternative family, those who have experienced a single-parent family only, and those who have had a step parent.[6] Consistent with the finding that daughters have a greater likelihood of experiencing divorce (Dahl & Moretti, 2004; Mammen, 2002; Morgan, Lye, & Condran, 1988), the proportion of those who have ever lived in a single or step parent family in the daughter sample is fifty percent higher than that of the son sample.

Although intergenerational mobility has been studied with respect to many economic outcomes including income, earnings, wealth, and occupation, it implicitly refers to some form of lifetime, permanent economic status. I examine the effects of family structure on three measures of labor market productivity – earnings, hourly wage, and education – for two reasons. First, measures

[4]Although the PSID does not provide a family structure history for each child, the information can be constructed using the marriage records of the parents. Thus, to be included in the samples used in this paper, a child must have non-missing identification information on both parents and non-missing marriage data on at least one parent. According to Bumpass and Raley (1995), definitions of single-parent families must be based on living arrangements rather than on marital status. Thus, every effort is made to identify cohabiting couples and assign them to the step family category. For instance, a household is believed to involve a cohabiting couple if family type determined by the marital status of the mother is single, but the head of the household is male. Blended families are not considered separately in this analysis. That is, one family may include a step child of one parent and a biological child of both parents. These half-siblings are no different in this analysis than whole siblings where the divorce occurred after the oldest reached age 18 and a re-marriage took place before the youngest was age 18.

[5]A few studies find that length of time spent in a single-parent family has no significant effect on the likelihood of high school graduation (Wojtkiewicz, 1993; McLanahan & Sandefur, 1994) or the risk of premarital birth (Wu & Martinson, 1993; Wu, 1996) and suggest that the number of family structure transitions is a better indicator of the impact of family structure on children. In this analysis, I include a measure of step family experience which should capture transitions as well as exposure.

[6]Note that most children who have lived in a step family would have first experienced a single-parent family. Also note that in ninety percent of the alternative families, the single parent is a single mother and the step parent is a step father.

associated with individuals rather than household measures of status are necessary for estimating mobility with respect to both mothers and fathers, resident or non-resident. Second, earnings alone are not a perfect measure of earnings potential because they depend not just on productivity or human capital but also on labor supply choices. Of particular concern for this analysis, family structure is related to labor force participation. Single mothers are more likely to work and given that they work, work longer hours on average (Employment Characteristics of Families, 2002). This is reflected in Table 1. Mothers have the highest average earnings when they are single and the lowest average earnings when they are married to the biological father. In contrast, they have the highest average wage and education (except in the son sample) in intact families, which implies that single mothers work more. There is also some evidence that men's hours may also be correlated to family structure. Sons with step parents and their fathers have higher average earnings than any other sons or fathers but have lower average hourly wages and educations than sons and fathers from intact families. Thus, sons with step parents and their fathers must work longer hours than the other groups. Because of these issues, using all three measures should increase the reliability of the findings.

To minimize the bias derived from measurement error in the dependent variable, I proxy the parents' permanent status with the log of an average of several years of earnings or several years of hourly wage observations, following Solon (1992) and Zimmerman (1992). Unlike Solon (1992) and Zimmerman (1992) but instead following Reville (1995), I also proxy the child's permanent status with an average of several years of observations to reduce the variance. The parent's earnings and hourly wage are taken from the 1968 through 1978 interviews and the child's adult earnings and hourly wage are taken from the 1983 to the 1993 interviews.[7] In Table 1, the numbers in square

[7] To be included in the sample, the child and both parents must have at least three years (and a maximum of eleven years) of self-reported earnings in which his/her employment status is not retired, disabled, a student, or other. In addition, the child must have been out of school for at least three years when the earnings are observed and less than age 19 in 1968 (the beginning of the panel). Annual earnings are defined in this paper as all wages and salaries including overtime, tips, commissions, bonuses, and any other form of payment for labor services received last year. Zero-earnings observations are included. Any observations

brackets are the average number of annual earnings observations used to construct the permanent earnings proxy. This number is lower for fathers and children from alternative families than for those from intact families, which implies greater measurement error among alternative families.[8] However, when I restrict the maximum number of observations such that the difference between intact and alternative families is small, I find no important difference in the results that follow. The differential measurement error by family structure occurs because the samples are restricted to biological parents and there are almost no observations for biological fathers who do not reside in the survey household of the mother and the children. Thus, the father's earnings and wage rate can only be observed before a divorce. In contrast, mother's earnings and hourly wage may represent post-divorce economic status.[9] However, if mothers' earnings are restricted to pre-divorce earnings, the results that follow are not affected.

I also provide the children's and the fathers' average ages by family structure to demonstrate that 1) the difference in the average age when earnings are observed for the child and the parent is ten or more years and 2) that the alternative families are slightly younger on average. Because of these factors, I control for the child's age in all regressions and the father's age in the average earnings and average wage regressions.[10]

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which have relevant variables imputed by 'major assignment' are excluded from the sample. The employment status of wives were not recorded until 1975 hence I do not restrict mothers by employment status.

[8]75% of intact fathers, but only 53% of divorced fathers, were observed for 8 or more years. At the other end of the spectrum, 4% of intact fathers, while more than 18% of divorced fathers, were observed for only 3 years.

[9]Since mothers are more likely to work after a divorce, divorced mothers have more annual earnings observations than intact mothers on average. 77% of intact mothers and 97% of divorced mothers were observed for 8 or more years. At the other end of the spectrum, 3% of intact mothers and no divorced mothers were observed in 3 years only.

[10]For the sake of parsimony, I do not control for mother's age.

3 The Effect of Family Structure on Earnings Mobility

3.1 The Model

The following framework motivates the estimation strategy presented in this paper. This model is derived from Becker and Tomes (1979) modified to allow for alternative families and for a difference in the impact of the earnings of mothers and fathers. Let us assume that a family involves a mother, a father, and one child, where either parent can be non-resident.

The family must allocate the sum of the parents' lifetime earnings $W_{t-1}^m + W_{t-1}^f$ between the family's current consumption C_{t-1} and investment I_{t-1} in the child's earning capacity, which gives us the following budget constraint:

$$W_{t-1}^m + W_{t-1}^f = C_{t-1} + I_{t-1}. \tag{1}$$

The child's earnings W_t are determined by the following function:

$$\begin{aligned} W_t &= (1+r)I_{t-1} \\ &+ \theta_m(1+\delta_m S_{t-1})W_{t-1}^m \\ &+ \theta_f(1+\delta_f S_{t-1})W_{t-1}^f \\ &+ L_t, \end{aligned} \tag{2}$$

where r is the return to human capital investment. θ_m and θ_f represent the degree to which the child inherits endowments from his mother and father, respectively, which allow the child to convert their earnings capacity into his own earnings capacity. The θ's can also be thought to represent the degree to which the child models his earnings-enhancing behavior based on his mother's and father's earnings. Thus, unlike in Becker and Tomes (1979), the earnings of parents can affect the child's earnings not only through the investment of their money, but also through role-modelling. Without these terms in the child's earnings production function, we would expect one dollar from the mother's earnings to have the same effect as one dollar from the father's earnings, which is contrary to findings in the literature. In particular, Couch and Dunn (1997) find that the earnings correlation with respect to mothers is lower than that with respect to fathers.

S_{t-1} represents whether (or how many years) the family has been a single or step parent family and the parameters δ_m and δ_f indicate the degree to which growing up in a single or step parent family affects the child's role-modelling or his ability to inherit endowments. Finally, L_t represents the child's 'market luck', which is assumed to be independent of W^m_{t-1} and W^f_{t-1} and known at the time of the investment.

The family chooses I_{t-1} to maximize the Cobb-Douglas utility function:

$$U = (1 - \alpha - \gamma S_{t-1}) \ln C_{t-1} + (\alpha + \gamma S_{t-1}) \ln W_t, \tag{3}$$

where the parameter α indicates the parents' average taste for W_t relative to C_{t-1}. The parameter γ indicates the degree to which the taste for investing in the child may differ for families involving non-resident fathers or single or remarried mothers.

Thus, this model allows alternative families to be different from intact families in two ways: 1) the child may be more or less able to inherit the endowments (or role-model) depending on δ_m and δ_f, and 2) the family may be more or less likely to invest in the child depending on γ.[11] The parameters can represent a causal link between family structure and child outcomes or a link caused by selection. For example, if $\gamma < 0$, we are not able to determine whether the parents invest less because of their family structure, or their family structure is merely an indicator of parents who invest less. The parameters δ_p where $p = \{m, f\}$ and γ can be positive or negative, however, at least one of them is negative since we know from the literature that divorce is correlated with negative outcomes (Amato & Keith, 1991; McLanahan, 1997; Grogger & Ronan, 1995; and Ginther & Pollak, 2000).

This model allows the earnings of parents to have different effects on daughters versus sons

[11] Note that for simplicity there is no mechanism in this model for the child to be affected by characteristics of the step parent, although that may be the case in reality.

given that there has been found to be a gender difference in mobility (Peters, 1992). In addition, there is evidence that alternative family experience affects children differentially by gender (Powell & Parcel, 1997). In particular, family structure may affect same-sex role-modelling through the role-modelling terms in the child's production function. For example, divorce may create a bond between mothers and daughters which causes the daughter to emulate the mother. Another example might be that step fathers have more influence on boys than on girls, which mitigates the mother's influence on her son after remarriage.

The first-order conditions imply that the optimal choice of I_{t-1} is:

$$\begin{aligned} I_{t-1} &= (\alpha + \gamma S_{t-1})(W_{t-1}^m + W_{t-1}^f) \\ &- \frac{(1-\alpha-\gamma S_{t-1})}{(1+r)}[\theta_m(1+\delta_m S_{t-1})W_{t-1}^m + \theta_f(1+\delta_f S_{t-1})W_{t-1}^f + L_t]. \end{aligned} \tag{4}$$

Substituting the optimal I_{t-1} into equation (2) produces:

$$\begin{aligned} W_t &= (\alpha + \gamma S_{t-1})(1+r)(W_{t-1}^m + W_{t-1}^f) \\ &+ (\alpha + \gamma S_{t-1})[\theta_m(1+\delta_m S_{t-1})W_{t-1}^m + \theta_f(1+\delta_f S_{t-1})W_{t-1}^f + L_t]. \end{aligned} \tag{5}$$

Re-writing the above equation yields:

$$\begin{aligned} W_t &= \beta_1 W_{t-1}^f + \beta_2 W_{t-1}^f S_{t-1} + \beta_3 W_{t-1}^f S_{t-1}^2 \\ &+ \beta_4 W_{t-1}^m + \beta_5 W_{t-1}^m S_{t-1} + \beta_6 W_{t-1}^m S_{t-1}^2 \\ &+ \beta_7 S_{t-1} + \beta_0 + u_t, \end{aligned} \tag{6}$$

where

$$\begin{aligned} \beta_1 &= \alpha(1+r) + \alpha\theta_f & \beta_4 &= \alpha(1+r) + \alpha\theta_m \\ \beta_2 &= \gamma(1+r) + (\gamma + \alpha\delta_f)\theta_f & \beta_5 &= \gamma(1+r) + (\gamma + \alpha\delta_m)\theta_m \\ \beta_3 &= \gamma\delta_f\theta_f & \beta_6 &= \gamma\delta_m\theta_m \end{aligned}$$

$$\begin{aligned} \beta_7 &= \gamma L_t \\ \beta_0 + u_t &= \alpha L_t. \end{aligned}$$

β_2 and β_3 represent the degree to which family structure affects the intergenerational relationship between children and fathers and β_5 and β_6 represent the effect of family structure on mother-child intergenerational earnings mobility.[12] If δ_f (in the case of β_2) and γ are of opposite sign, the model does not allow us to predict the sign of β_2.[13] That is, we cannot predict whether mobility with respect to either parent will be higher or lower in alternative families compared to intact families.

3.2 Empirical Strategy

Bearing this model in mind, we can study intergenerational earnings mobility and its interaction with family structure using the following regression equation:

$$\begin{aligned} W^c_{ji} &= \rho_1 W^f_j + \rho_2 W^f_j s_{ji} + \rho_3 W^f_j s^2_{ji} + \rho_4 W^f_j step_{ji} \\ &+ \rho_5 W^m_j + \rho_6 W^m_j s_{ji} + \rho_7 W^m_j s^2_{ji} + \rho_8 W^m_j step_{ji} \\ &+ \rho_9 s_{ji} + \rho_{10} s^2_{ji} + \rho_{11} step_{ji} + \rho_{12} A_i + \rho_0 + \varepsilon_{ji}, \end{aligned} \tag{7}$$

The logarithm of the permanent earnings proxy for child i in family j (W^c_{ji}) is a linear function of the logarithm of the permanent earnings proxies of each parent (W^f_j and W^m_j), his or her exposure to a single or step parent family (s_{ji}), interactions between parents' earnings and duration in an alternative family, and a vector (A_i) of the child's median age (and age squared) when his/her earnings were observed, the father's median age (and age squared) when his earnings were observed, and interactions between the family structure variables and the father's age.[14] This specification allows for the number of years in a single or step parent family to affect mobility non-linearly and the number of years in a step family ($step_{ji}$) to have a differential effect since children in step

[12] Note that the interaction in the third and sixth terms of equation (6) are not distinct from that in the second and fifth terms, and hence redundant, if S_{t-1} is a dummy variable which only takes the values zero and one.

[13] Likewise, if δ_m and γ are of opposite sign, we cannot predict the sign of β_5 given this model.

[14] For the sake of parsimony, I do not control for mother's age.

families have likely experienced more transitions in family structure and there is the involvement of another parent figure. This equation is identical to equation (6) from the model with the addition of the step family indicator and the age controls.

Both mother's and father's earnings are included in the regression because running separate regressions would ignore assortative mating. That is, if mother's earnings are merely an indicator of father's earnings, then the importance of mother's earnings will be exaggerated in a specification that does not control for father's earnings. Finally, it is assumed that $E(\varepsilon_{ji} \mid z_{ji}) = 0$ and $V(\varepsilon_{ji} \mid z_{ji}) = \sigma^2\Omega$, where z_{ji} is the vector of regressors listed in equation (7), the diagonal elements of Ω are ω_i^2, the off-diagonal elements within families are $\omega_{ij}\omega_{ji}$, and the off-diagonal elements across families are zero (i.e. the standard errors are adjusted for intra-cluster correlations at the family level).

I estimate the coefficients using ordinary least squares. If ρ_2 is negative (and assuming that $\rho_3 = 0$ and $\rho_4 = 0$), then each additional year living in an alternative family is associated with a level of intergenerational earnings mobility with respect to the father that is ρ_2 higher.

3.3 Results

3.3.1 Sons

Table 2 presents the mobility of sons with respect to their fathers and mothers using log average earnings as the proxy for permanent earnings. The coefficient in the top left-hand corner indicates that the elasticity of sons' earnings with respect to father's earnings is 0.299.[15] The bottom half of Table 2 presents the coefficients on mother's log earnings and the family structure interactions. The coefficient on mother's log average earnings indicates that the elasticity of son's earnings with respect to mother's earnings, holding father's earnings constant, is positive and significant but only

[15] The magnitude of this coefficient is smaller than the 0.4 estimate found by Solon (1992) and Zimmerman (1992) because it involves a broader cohort. That is, both Solon and Zimmerman use son's earnings from the 1985 interview where this sample involves son's earnings from the 1983 through the 1993 interviews. When I restrict the son's earnings to just those earnings observed in 1985, the elasticity of son's earnings with respect to father's earnings is 0.464, which is more in line with the findings in the literature.

five percent of the father's coefficient in size.

The coefficient at the top of column (2) indicates that, for sons with zero years in a single or step parent family, the earnings elasticity with respect to father's earnings is 0.315. In this column, I include a variable indicating whether the son ever lived in an alternative family (the main effect coefficient is not shown) and interactions between this variable and the parents' log average earnings. The coefficient on the father's interaction is negative and significant indicating that sons with experience in an alternative family are more mobile with respect to their fathers than sons who have only lived in a intact family. On the other hand, the coefficient on the mother's interaction is not significantly different from zero.

The family structure variable used in columns (3) through (5) is the number of years spent in either a single or step parent family. Column (5) also includes a measure of the number of those years which were in a step parent family. The different specifications reveal that the effect of duration on father-son mobility is non-linear. That is, the effect on father-son mobility of spending one year in an alternative family relative to zero years is much larger than the effect of spending 10 years relative to 9 years. The coefficients on the mother's interactions are not significantly different from zero, however the coefficients on the interactions between father's and mother's log earnings and years in a step family are negative, large, and significant. Thus, the type of alternative family has an effect on son's mobility with respect to both his mother and father as we would expect given that children in step families have experienced more transitions in family structure than those in single parent families only and there is the involvement of another parent figure.

Based on the coefficients from the specification in column (5), a son with three and a half years in single-parent family – the average for the son and daughter samples – has an intergenerational elasticity with respect to his father or his mother that is not significantly different from zero. On the other hand, a son with three years in a step parent family on top of three and a half years in a single parent family has an intergenerational elasticity with respect to his father and his mother that is negative and significant at the 10% level. A possible explanation for the negative imputed father-son elasticity is that, among sons with step parent family experience, those with the highest earning biological fathers end up with the worst outcomes. This is consistent with the hypothesis that the transitions in family structure matter in that these sons would have had more transitions and potentially a dramatic transition if the fathers' high earnings are an indication of the fall in household income at the time of divorce. The negative mother-son elasticity may be explained by a time investment story. That is, if a mother remarries and works long hours so that her earnings are high, the time she has available to invest in her son is limited and his future earnings suffer as a result.

3.3.2 Daughters

Table 3 presents the mobility of daughters with respect to their fathers and mothers. Column (1) indicates that the daughters' earnings elasticity with respect to their fathers' earnings is 0.399.[16] Unlike the sons sample, the coefficient on mother's log average earnings indicates that the elasticity of daughter's earnings with respect to mother's earnings, holding father's earnings constant, is zero. Although it is not intuitive that mother's earnings matter more to sons than daughters, the finding is consistent with the fact that McLanahan (1985) finds no support for same-sex role modelling.

In column (2), the coefficient on the interaction between the father's earnings and the indicator of ever having lived in an alternative family is negative and significant indicating that daughters with experience in an alternative family are more mobile with respect to their fathers than daughters who have only lived in a intact family. In contrast, the coefficient on the interaction between mother's earnings and alternative family experience is positive and significant indicating that there is less mother-daughter mobility among alternative families. Thus, although daughters do not appear

[16] The magnitude of this coefficient is larger than that found in the small literature which analyzes father-daughter earnings mobility; Altonji and Dunn (1991) estimate an elasticity of 0.22 and Peters (1992) estimates an elasticity of 0.11. However, both of these studies use the NLS instead of the PSID, which can explain the difference. The PSID has been used in a study of daughter mobility with respect to *family income* (Minicozzi, 1997) with an estimate of 0.41, which is more in line with the estimate presented here.

to be modelling their mothers in intact families, mother-daughter role modelling appears to be important in alternative families.

Similar to the father-son mobility findings, the coefficients in columns (3) through (5) suggest that the effect of the number of years in an alternative family on both father-daughter and mother-daughter mobility is non-linear. The type of alternative family has a significant effect on father-daughter mobility but not on mother-daughter mobility.

Based on the coefficients from the specification in column (5), a daughter with three and a half years in single-parent family has an intergenerational elasticity with respect to her father that is not significantly different from zero and an elasticity with respect to her mother of 0.177 (significant at the 1% level). On the other hand, a daughter with three years in a step parent family and three and a half years in a single parent family has an intergenerational elasticity with respect to her father that is negative and significant at the 10% level and an elasticity with respect to her mother that is not significantly different from zero. The imputed elasticities for daughters given step family experience follow the same pattern as the corresponding imputed elasticities for sons. That is, the negative father-daughter elasticity with step family experience is consistent with the hypothesis regarding the importance of transitions and the insignificant mother-daughter elasticity with step family experience is consistent with the time investment story given that the mother-daughter elasticity is positive without the step family experience.

These findings are consistent with those of Biblarz and Raftery (1993) who examine the effect of family structure on intergenerational *occupational* mobility using data from the 1973 Occupational Changes in a Generation Survey and find that father-son mobility is higher for alternative families. In contrast, Peters (1992) uses data from the National Longitudinal Survey (NLS) and finds that the family characteristic 'broken home' does not have a significant effect on intergenerational earnings mobility of sons or daughters with respect to their parents. However, if I use the family structure variable in Peters (1992), a single snap-shot at age 14, instead of 'any exposure' or 'years of exposure' as I do in this study, I also find an insignificant effect on mobility.

4 The Mobility of Hourly Wage and Education

The results in the previous section indicate that father-son and father-daughter mobility are higher and mother-daughter mobility is lower in alternative families. However, the specification used above (laid out in equation (7)) is not without problems. First, there is measurement error in the proxies for the parents' permanent earnings which biases the coefficients toward zero. Second, mother's earnings, even an average of several years, is not a good measure of mother's earnings potential given the uneven patterns of women's labor supply. In fact, single mothers may have higher earnings, as we observe in Table 1, only because their work hours are greater, not because they have higher earnings potential. This problem also results in attenuation given that this earnings proxy will confound mothers with valid estimates of their permanent earnings with mothers who work irregular hours or who do not work every year.

To address these issues, I use two alternate measures of labor market productivity: hourly wage and education.[17] Education suffers from much less measurement error than average earnings given that it is easy to quantify and recall. It also is not subject to issues related to labor supply. The latter argument is the main advantage of using hourly wage. When the hourly wage of the parents are used, the sample must be restricted to those who have some earnings. Thus, the hourly wage provides a measure of earnings independent of the mother's labor force participation and hours decisions.

Table 4 presents the mobility of sons and daughters with respect to their fathers and mothers using log average hourly wage and education as proxies for permanent earnings. I also include for comparison the coefficients from the regression using parents' log average earnings from column (5) of Tables 2 and 3.

[17] I do not control for the parent's age and age squared when using education as I do for earnings and hourly wage since the age-education profile is relatively flat by the mid-20s.

For the son's sample, the coefficients on father's hourly wage and the father's wage interactions in column (2) are similar to those on father's earnings and the corresponding interactions in column (1). Some of the estimates lose significance because the sample is smaller in the wage regressions. The coefficients on mother's hourly wage and, in particular, the coefficients on the mother's wage-family structure interactions at the bottom of column (2) are different from the corresponding coefficients in the earnings regression at the bottom of column (1), however they are not significant. In sum, this specification provides some support for the claims that father-son mobility is higher and that mother-son mobility is the same in alternative families.

For the daughter's sample, the coefficients on the mother's hourly wage and wage interactions in column (5) are similar in magnitude and sign to those in the earnings regression in column (4), but the coefficients on the father's wage interactions are different from those in the earnings regressions. The similarity between mothers and daughters may stem from the relative similarity in a mother's and daughter's labor supply choices. The difference in the father's interaction coefficients may be due to a combination of less variability in daughters' hourly wage compared to their earnings and a lack of precision in the wage regression.

The coefficients in the education regressions are consistent in general with the findings in the other columns with some specific differences. First, mother's education has a significant positive effect on both son's and daughter's earnings, and the mothers' coefficients are half the size of the fathers' coefficients, unlike the relatively small impact of mother's earnings and hourly wage compared with fathers'. Second, the effect of family structure on mobility appears to be significant only for those with step family experience. That is, both sons and daughters living in step parent families have higher mobility with respect to their fathers and lower mobility with respect to their mothers and the differences are mostly significant (not significantly for sons).

Overall, these alternate specifications provide some support for the claim that father-child mobility is higher and mother-daughter mobility is lower in single and step parent families. Thus, my interpretation of these alternate specifications is that the earnings regressions are not misleading despite the fact that earnings, particularly mothers' earnings, are a flawed measure of economic status.

5 Causality

Given that I find a relationship between family structure and mobility, is it causal? That is, is it another consequence of divorce that children are less like their fathers and more like their mothers in terms of economic status, or is it the case that families who have high father-child and low mother-child mobility are more likely to experience divorce? I address this issue in two ways. First, following Grogger and Ronan (1995), and Case, Lin, and McLanahan (2000), I look at sibling differences in alternative family experience using fixed effects (FE).[18] Because this technique has an important shortcoming which I describe below, I also conduct an alternate analysis which I describe later.

5.1 Sibling Fixed Effects

Let φ_f be the sibling fixed effect. I then estimate the following equation:

$$\begin{aligned} W^c_{ji} &= \rho_2 W^f_j s_{ji} + \rho_3 W^f_j s^2_{ji} + \rho_4 W^f_j step_{ji} \\ &+ \rho_6 W^m_j s_{ji} + \rho_7 W^m_j s^2_{ji} + \rho_8 W^m_j step_{ji} \\ &+ \rho_9 s_{ji} + \rho_{10} s^2_{ji} + \rho_{11} step_{ji} + \rho_{12} A_i + \varphi_f + \varepsilon_{ji}, \end{aligned} \tag{8}$$

where s_{ji} represents the number of years in a single or step parent family and A_i is a vector of the child's median age (and age squared) when his/her earnings were observed and interactions between the family structure variables and the father's age. This strategy identifies the effect of family structure on intergenerational earnings mobility from the differences in the duration of spells

[18] I estimate the fixed effects in the regressions presented here, although I have also estimated the coefficients using sibling differences and have gotten similar estimates.

living in an alternative family structure across siblings. The fixed effect should capture the time-invariant unobservable characteristics of a family. If the family selected into their family structure based on any of these unobservable characteristics, the estimated effect of family structure should not be driven by that type of selection.

Table 5 provides some statistics on differences across siblings since the sibling fixed effects analysis hinges on the existence of these differences. Because of the age differences between siblings, children in alternative families will have spent different numbers of years in each family type. Of those whose families were alternative at some time, on average the difference spent in alternative families across siblings is between 3 and 4 years. This specification assumes that the impact of alternative families increases with exposure. Because duration in an alternative family type is proportional to the age at disruption, this assumption could lead to a misinterpretation of the results if there were other consequences of divorce related to age at disruption or strong birth order effects. However, I test for the presence of either of these factors and find that they do not hold.[19]

There are two other noteworthy characteristics of the sibling sample. First, the alternative families have a higher proportion of daughters because parents with a son are less likely to divorce (Dahl & Moretti, 2004; Mammen, 2002; Morgan et al., 1988). Finally, because the sibling sample does not include 'only children,' which make up a large proportion of single-mother households, the family size statistics indicate that alternative families are smaller on average. That is, the intact families are more evenly distributed across one, two, and three or more siblings where the alternative families are highly concentrated among the one-sibling families.

[19]The 'family stress' hypothesis claims that the impact of divorce on young children involves behavioral problems, where older children experience a withdrawal from the family. If this hypothesis holds, the mobility of children with shorter durations in alternative families could in principle be more affected by family structure than children with longer durations. To test this possibility, I separate children by whether the divorce occurred before or after age 10. I find that the impact is significant only for children who experience the disruption at a young age, indicating that the data do not support the 'family stress' hypothesis. If there were birth order effects, lower birth order children, who would have had more limited experience with alternative families, would always appear to do better than higher birth order children making the effect of alternative families appear worse. A test on children within intact families only indicates that older children do not have any significant advantage over younger children.

Table 6 provides the results from the fixed effects regression. Because the identification strategy is based on the variation across siblings in the number of years in an alternative family, only the duration measure of family structure is used in this analysis. The fixed effects regressions cannot include the main effect of father's or mother's earnings because it does not vary across siblings.[20] I include the results from a standard OLS regression to demonstrate that the estimated effect of including the fixed effect is not the result of changing from a single-sex sample to the mixed-sex sibling sample. As in the single-sex sample regressions, father-child mobility is higher in single and step parent families, and step parent families particularly increase both father-child and mother-child mobility. Because there are more daughters among those with alternative family experience, in the sibling sample mother-child mobility is lower in alternative families, as is the case using the daughters sample.

In the FE model, the coefficient on the interaction between father's earnings and alternative family experience drops to a tenth the size of the OLS coefficient and is not significantly different from zero.[21] The FE coefficient on the interaction between father's earnings and step family experience is not statistically different from the OLS coefficients. Thus, father-son mobility is the same for siblings who spent different spells in single parent families, but higher for the sibling who spent more time in a step parent family. This implies that the association between any alternative family experience and father-child mobility is driven by selection on time-invariant family unobservable characteristics. This finding is not consistent with the idea that time away from the father is the cause of the higher mobility. However, the step family interaction suggests that time with a step father, or the transition into remarriage, may have a causal effect on father-son mobility.

[20] Father's age also does not vary across siblings and so the fixed effects specifications do not include father's age as a control.

[21] However, the difference between the FE and the OLS coefficient is not significant. The statistical significance of the difference between the coefficients was estimated by bootstrapping the standard errors and estimating the covariance between the βs. In row (2), the difference between columns (1) and (2) is 0.416 (0.510).

In contrast, the FE coefficient on the interaction between mother's earnings and alternative family experience is not statistically different from the OLS coefficients but the difference between the FE and OLS coefficients on the interaction between mother's earnings and step family experience is significant. Thus, mother-child mobility is lower for siblings who have spent more time in an alternative family and the effect is larger for those with step family experience, which implies that the correlation between family structure and mother-child mobility is not the result of this type of selection.[22] Since this does not rule out a causal explanation, it may be the case that increasing a child's time with a single or remarried mother strengthens their earnings correlation.

There is one important potential problem with interpreting the sibling fixed effects coefficients as causal, particularly the coefficient on the father's interaction. Because I find the effect of years in an alternative family to be non-linear in Tables 2 and 3, particularly with respect to fathers, the fixed effects coefficients may be downward biased. That is, the effect of divorce on mobility is greatest in the first year after the divorce and has smaller incremental effects in the years following (i.e. the coefficient on parent's earnings interacted with *years squared* is significant and the opposite sign of the coefficient on parent's earnings interacted with *years*). Thus, the difference in the effects across siblings should be small because most of the effect occurs for both siblings. The effect of family structure on mobility is more non-linear with respect to fathers than mothers, which implies that the bias should be greater for fathers. This bias may be responsible for the small and insignificant coefficient on the father's interaction, rather than the selection of families into divorce.

[22] If daughters are likely to be the youngest sibling in divorced families with different sex siblings, then the coefficient on the mother's interaction with family structure might be biased upward. Because daughters are more like their divorced mothers than sons (from Tables 2 and 3), the fixed effects specification may be picking up this gender difference rather than the age difference on which I argue identification is based. However, the youngest child is a daughter in only 41.4% of divorced families with different sex siblings. In fact, the oldest child is a daughter in 65.5% of these families, which suggests that, if anything, the coefficient on the mother's interaction may be biased downward.

5.2 Instrumental Variables

Because of the potential downward bias of the sibling fixed effect technique, I also use instrumental variables (IV) to investigate the causality of the relationship between family structure and mobility. In particular, I instrument the number of years in a single or step parent family with the child's exposure to no-fault divorce laws. The number of divorces increased following the passing of no-fault divorce laws (Friedberg, 1998) and thus the number of years the child spent in a state which permitted no-fault divorces serves as exogenous variation in the probability that his/her parents' divorce.

No fault divorce increased the ease of divorce by not requiring that one spouse demonstrate a transgression by the other; instead, irreconcilable differences could be claimed as grounds for divorce. According to the legislative details compiled in Gruber (2000), there were a dozen states with no-fault divorce laws in place by 1950. A dozen more passed such laws in the 1950s and 1960s, but the majority were passed in the 1970s, such that by 1985, all states permitted no-fault divorce. Many states also simultaneously or subsequently permitted unilateral divorce, which increased the ease of divorce by not requiring the consent of both spouses.

Several recent studies have examined changes in unilateral divorce laws (Friedberg, 1998; Gruber, 2000; Wolfers, 2003) but I focus on no-fault divorce as the timing of these laws are more relevant to the cohorts of children in this sample of the PSID (born between 1948 and 1968). Table 7 presents some summary statistics regarding the children's exposure to no-fault and unilateral divorce laws by family structure. For all family structure types, the mean number of years exposed to no-fault is at least 4 years greater than the mean number of years exposed to unilateral divorce. In addition, among intact families, the fraction of children who were never exposed to no-fault divorce is half that of unilateral divorce. Stikingly, more than half of the children from divorced families were never exposed to unilateral divorce laws. Because the variation in exposure to unilateral divorce is so low for this sample, I look only at no-fault divorce laws.

As expected given the literature on the effects of divorce laws on divorce, the mean exposure is greater for those children whose parents divorced. Further, the mean exposure is higher for those children who have experienced a step family compared to those who have not which is consistent with the fact that these children have longer average durations in alterative families.

Table 8 presents the results from this IV estimation procedure. I do not present the first stage coefficients but they behave appropriately and the regression has an F-statistic of 16.17. I include the results from an OLS regression for comparison. As above, the coefficients in the OLS model using this pooled sample of boys and girls are similar to those found with the single-sex samples.

The coefficient sizes and signs on the father's earnings and particularly the interactions, are strikingly similar across the two regressions presented here. However, as is the usual case with IV models, the standard errors are much larger than in the OLS model. As a result, only the main effect of father's earnings is significant in the IV model. In contrast, the coefficient sizes on mother's earnings and its interactions are different but the signs are the same and the significance is more in line with the OLS model. In particular, the interaction between mother's earnings and alternative family experience is positive and significant, consistent with both the OLS coefficient and the sibling fixed effects coefficient.

Thus, the IV analysis confirms the causal interpretation of the mothers' results – mother-child mobility falls with each additional year in an alternative family. It also confirms, although weakly given the similar coefficient size on the father's interaction, that father-child mobility appears higher in alternative families because of selection.

6 Family Structure and Mobility Trends

In the sections above, I show that there is a relationship between family structure and father-child mobility, although it is likely not causal. In this section, I examine whether this relationship taken together with the rising prevalence of alternative families in recent decades is responsible for rising mobility between fathers and sons. The relationship need not be causal for rising mobility to occur.

That is, father-son mobility may be higher because more fathers spend less time with their sons because of a divorce, or because some fathers have changed – they both have a smaller influence on their sons and select into divorce more.

To test the effect of changing family structure on the trend in mobility, I construct a father-son sample with no siblings (for consistency with the mobility trends literature) and divide it into five cohorts. The results of this analysis are reported in Table 9. In the first column, I report the proportion of the sample which spent any time in an alternative family. Given the trend in the US, this proportion increases across the cohorts. In the top panel, I estimate the father-son intergenerational elasticity for each of the five cohorts; in column (2), I do not control for family structure and its interactions with father's earnings, while in column (3) I do. In the first cohort, the intergenerational elasticity is approximately 0.5 in both columns. Starting in the second cohort, the intergenerational elasticity falls more dramatically in the column with no controls than in the other column. The bottom panel displays the results of pooling all of the cohorts and including a cohort indicator and an interaction between cohort and father's earnings. The decline in the intergenerational elasticity – represented by the coefficient on the cohort interaction – is significant only in the column with no family controls and the difference in the coefficients on the cohort interaction variable across the specifications is significant. This evidence suggests that changes in family structure may be responsible for half, if not all, of the rise in father-son mobility in the US in recent decades.

7 Summary and Conclusion

In this study, I use the PSID to investigate the association between family structure and intergenerational earnings mobility. I find that with each additional year in either a single or step parent family, the mobility of biological sons and daughters are higher with respect to their father, and the mobility of daughters is lower with respect to their mother. These findings hold up when the parents' economic status is proxied with their hourly wage and education. I also find that the association between any alternative family experience and father-child mobility is explained by selection on family unobservable characteristics where the correlation between family structure and mother-child mobility is not.

These findings do not offer support for the father-absence hypothesis since it appears that the connection between fathers and children would have been weak whether or not a divorce occurred. These findings also indicate that the overall level of father-son mobility has been rising because there are more alternative families in recent cohorts. Averaging the mobility levels of both intact families and alternative families make it appear that overall mobility is rising when in reality the mobility levels of each type of family is not changing over the period. Taken together, these results suggest that it is not divorce, or father absence in particular, that affects father-child mobility but rather that fathers have changed in recent decades. There are more fathers who are both divorcing and having a reduced impact on their children's outcomes. Normatively, since poorer couples are more likely to divorce, higher mobility among single and step parent families suggests that these children are not destined to be as poor as their fathers. That is, if fathers have changed for the worse, it is to the benefit of their children to have higher mobility.

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Table 1: Summary Statistics

	Sons Sample			Daughters Sample		
	Always Intact	Only Single	Any Step	Always Intact	Only Single	Any Step
Percent	93.7	4.0	2.3	90.4	5.5	4.2
Mean Years		3.7	5.9		3.5	7.3
if in Category		(0.5)	(0.7)		(0.4)	(0.5)
Mean Child's						
Earnings	$23,057	$22,435	$23,775	$22,685	$19,951	$15,759
	(702)	(2,128)	(3,694)	(612)	(1,930)	(1,716)
	[8.6]	[7.0]	[6.8]	[8.8]	[7.2]	[7.6]
Wages	$10.72	$10.01	$10.06	$10.83	$9.46	$8.51
	(0.28)	(0.84)	(1.19)	(0.26)	(0.86)	(0.63)
Years of Schooling	12.8	12.8	10.6	12.9	12.7	10.5
	(0.1)	(0.6)	(1.4)	(0.1)	(0.4)	(0.9)
Mean Father's						
Earnings	$31,542	$28,203	$35,432	$33,480	$27,908	$32,378
	(774)	(2,374)	(3,636)	(887)	(1804)	(2072)
	[9.2]	[6.9]	[6.6]	[9.0]	[7.3]	[6.2]
Wages	$13.99	$12.44	$13.69	$14.39	$12.25	$13.61
	(0.33)	(1.02)	(1.52)	(0.36)	(0.78)	(0.89)
Years of Schooling	12.7	12.1	12.4	12.9	12.1	12.8
	(0.1)	(0.4)	(0.6)	(0.1)	(0.3)	(0.5)
Mean Mother's						
Earnings	$4,400	$5,538	$5,252	$4,299	$6,503	$6,259
	(227)	(953)	(1259)	(216)	(971)	(974)
	[8.7]	[9.8]	[9.4]	[8.5]	[9.4]	[9.0]
Wages	$6.91	$5.84	$6.58	$7.18	$6.42	$6.85
	(0.16)	(0.42)	(0.51)	(0.26)	(0.42)	(0.24)
Years of Schooling	11.5	12.0	9.8	12.5	12.0	11.9
	(0.1)	(0.6)	(1.2)	(0.1)	(0.3)	(0.6)
Mean Child's Age	30.7	27.6	27.1	30.2	27.1	26.2
	(0.2)	(0.5)	(0.4)	(0.2)	(0.5)	(0.4)
Mean Father's Age	43.7	37.6	35.4	43.6	38.7	34.6
	(0.3)	(1.1)	(1.3)	(0.3)	(1.5)	(1.3)
Total Individuals		701			695	
Total Families		465			468	

Standard errors are in parentheses. The average number of annual earnings observations are in square brackets. The earnings figures are deflated using the CPI and presented in 1984 dollars.

Table 2: Sons Results

Dependent Variable: 3+ Year Average of Log Son's Earnings

	(1)	(2)	(3)	(4)	(5)
F's Log Earnings	0.299**	0.315**	0.303**	0.314**	0.315**
	(0.074)	(0.076)	(0.075)	(0.076)	(0.076)
F's Log Earnings		-0.295+			
* Alt. Family		(0.159)			
F's Log Earnings			-0.030	-0.234**	-0.255**
* Yrs in Alt. Family			(0.080)	(0.094)	(0.079)
F's Log Earnings				0.034+	0.030
* Yrs^2 in Alt. Family				(0.020)	(0.020)
F's Log Earnings					-1.402*
* Yrs in Step Family					(0.646)
M's Log Earnings	0.017+	0.016	0.015	0.016+	0.016
	(0.009)	(0.010)	(0.010)	(0.010)	(0.010)
M's Log Earnings		0.004			
* Alt. Family		(0.055)			
M's Log Earnings			0.008	-0.024	0.013
* Yrs in Alt. Family			(0.015)	(0.024)	(0.036)
M's Log Earnings				0.005	0.004
* Yrs^2 in Alt. Family				(0.004)	(0.004)
M's Log Earnings					-0.147*
* Yrs in Step Family					(0.065)
Sample Size			701		

Heteroskedasticity-rodust standard errors adjusted for intra-cluster correlations are in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. In all specifications, son's age, son's age squared, father's age, and father's age squared are included. In the specifications with interactions, the family structure variables and the interactions between the family structure variables and father's age and father's age squared are also included.

Table 3: Daughters Results

Dependent Variable: 3+ Year Average of Log Daughter's Earnings

	(1)	(2)	(3)	(4)	(5)
F's Log Earnings	0.399**	0.471**	0.442**	0.461**	0.465**
	(0.065)	(0.067)	(0.064)	(0.066)	(0.066)
F's Log Earnings * Alt. Family		-0.528*			
		(0.227)			
F's Log Earnings * Yrs in Alt. Family			-0.062	-0.244**	-0.276**
			(0.040)	(0.099)	(0.080)
F's Log Earnings * Yrs^2 in Alt. Family				0.016*	0.046**
				(0.008)	(0.017)
F's Log Earnings * Yrs in Step Family					-0.436*
					(0.236)
M's Log Earnings	0.002	-0.007	-0.002	-0.007	-0.006
	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)
M's Log Earnings * Alt. Family		0.165**			
		(0.069)			
M's Log Earnings * Yrs in Alt. Family			0.023**	0.148**	0.091**
			(0.009)	(0.053)	(0.031)
M's Log Earnings * Yrs^2 in Alt. Family				-0.019**	-0.011*
				(0.008)	(0.005)
M's Log Earnings * Yrs in Step Family					-0.044
					(0.028)
Sample Size			695		

Heteroskedasticity-rodust standard errors adjusted for intra-cluster correlations are in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. In all specifications, daughter's age, daughter's age squared, father's age, and father's age squared are included. In the specifications with interactions, the family structure variables and the interactions between the family structure variables and father's age and father's age squared are also included.

Table 4: Alternate Specifications
Dependent Variable: 3+ Year Average of Log Child's Earnings

	Son's Sample			Daughter's Sample		
Parents' Economic Status=	Avg Earnings	Avg Wages	Yrs of Schooling	Avg Earnings	Avg Wages	Yrs of Schooling
F's Economic Status	0.315**	0.434**	0.066**	0.465**	0.570**	0.088**
	(0.076)	(0.097)	(0.018)	(0.066)	(0.097)	(0.024)
F's Economic Status	-0.255**	-0.204	0.065	-0.276**	-0.086	0.011
* Yrs in Alt. Family	(0.079)	(0.199)	(0.047)	(0.080)	(0.194)	(0.039)
F's Economic Status	0.030	0.019	-0.001	0.046**	0.003	0.004
* Yrs^2 in Alt. Family	(0.020)	(0.032)	(0.006)	(0.017)	(0.026)	(0.003)
F's Economic Status	-1.402*	-0.462	-0.109**	-0.436*	0.076	-0.096*
* Yrs in Step Family	(0.646)	(0.362)	(0.039)	(0.236)	(0.342)	(0.043)
M's Economic Status	0.016	0.086	0.030*	-0.006	-0.071	0.043+
	(0.010)	(0.085)	(0.015)	(0.014)	(0.111)	(0.024)
M's Economic Status	0.013	-0.101	-0.023	0.091**	0.111	-0.036
* Yrs in Alt. Family	(0.036)	(0.208)	(0.044)	(0.031)	(0.262)	(0.037)
M's Economic Status	0.004	0.023	0.000	-0.011*	0.002	-0.001
* Yrs^2 in Alt. Family	(0.004)	(0.032)	(0.006)	(0.005)	(0.042)	(0.003)
M's Economic Status	-0.147*	0.327	0.018	-0.044	-0.142	0.114*
* Yrs in Step Family	(0.065)	(0.322)	(0.045)	(0.028)	(0.436)	(0.053)
Sample Size	701	569	689	695	574	685

Heteroskedasticity-rodust standard errors adjusted for intra-cluster correlations are in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. In all specifications, the family structure variables, child's age, and child's age squared are included. In the earnings and wage specifications, father's age, father's age squared, and the interactions between the family structure variables and father's age and father's age squared are also included.

Table 5: Summary Statistics on Sibling Sample

	Always Intact	Only Single	Any Step
% in Category	91.9	4.7	3.4
% Daughters	48.5	56.6	63.2
Mean Difference in Years		3.2	3.6
		(0.3)	(0.4)
Max Difference in Years		11	11
% with No Difference in Years		4.9	0.0
% with 1 sibling	34.7	56.6	39.5
% with 2 siblings	33.4	30.2	21.1
% with 3+ siblings	31.9	13.2	39.5
Total Individuals		1125	
Total Families		402	

Standard errors are in parentheses.

Table 6: Fixed Effects Results

Dependent Variable: 3+ Year Average of Log Child's Earnings

	OLS	FE
F's Log Earnings	0.393**	
	(0.065)	
F's Log Earnings * Yrs in Alt. Family	-0.379**	0.037
	(0.086)	(0.047)
F's Log Earnings * Yrs^2 in Alt. Family	0.068**	0.002
	(0.029)	(0.006)
F's Log Earnings * Yrs in Step Family	-0.483	-0.303**
	(0.501)	(0.098)
M's Log Earnings	0.004	
	(0.009)	
M's Log Earnings * Yrs in Alt. Family	0.079**	0.096**
	(0.028)	(0.034)
M's Log Earnings * Yrs^2 in Alt. Family	-0.006+	-1.290+
	(0.004)	(0.679)
M's Log Earnings * Yrs in Step Family	-0.108**	4.692+
	(0.043)	(2.613)
Sample Size	1125	1125

Heteroskedasticity-rodust standard errors adjusted for intra-cluster correlations are in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. In both specifications, child's age, child's age squared, child's sex, the family structure variables, and the interactions between the family structure variables and father's age and father's age squared are included. In the OLS specification, father's age and father's age squared are also included.

Table 7: Summary Statistics on Divorce Law Exposure

	Always Intact	Only Single	Any Step
Mean years of exposure to no-fault divorce laws	6.9	8.3	11.5
	(0.2)	(0.7)	(0.8)
% with No Exposure	30.0	9.1	0.0
Mean years of exposure to unilateral divorce laws	2.2	3.3	3.8
	(0.1)	(0.5)	(0.7)
% with No Exposure	66.8	53.0	55.6
Total Individuals		1396	
Total Families		673	

Standard errors are in parentheses.

Table 8: Instrumental Variables Results
Dependent Variable: 3+ Year Average of Log Child's Earnings

	OLS	IV
F's Log Earnings	0.386**	0.266*
	(0.058)	(0.132)
F's Log Earnings * Yrs in Alt. Family	-0.295**	-0.305
	(0.063)	(0.443)
F's Log Earnings * Yrs^2 in Alt. Family	0.043**	-0.040
	(0.015)	(0.088)
F's Log Earnings * Yrs in Step Family	-0.395	-0.399
	(0.262)	(0.397)
M's Log Earnings	0.005	-0.027
	(0.009)	(0.034)
M's Log Earnings * Yrs in Alt. Family	0.058**	0.873*
	(0.025)	(0.425)
M's Log Earnings * Yrs^2 in Alt. Family	-0.005	-0.115**
	(0.003)	(0.047)
M's Log Earnings * Yrs in Step Family	-0.075**	-0.749
	(0.030)	(0.861)
Sample Size	1396	1396

Heteroskedasticity-rodust standard errors adjusted for intra-cluster correlations are in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. In both specifications, child's age, child's age squared, child's sex, the family structure variables, father's age, father's age squared, and the interactions between the family structure variables and father's age and father's age squared are included.

Table 9: Father-Son Trends
Elasticity of Son's Earnings w.r.t Father's Earnings
Dependent Variable: 3+ Year Average of Log Child's Earnings

	Proportion of Sample in Alt. Families	No Family Controls	Family Controls	Difference
'68-'85 Cohort	3.1	0.497**	0.485**	
[384]		(0.101)	(0.104)	
'69-'86 Cohort	4.4	0.320**	0.359**	
[389]		(0.081)	(0.084)	
'70-'87 Cohort	4.4	0.184*	0.295**	
[385]		(0.080)	(0.081)	
'71-'88 Cohort	4.6	0.202**	0.373**	
[393]		(0.081)	(0.085)	
'72-'89 Cohort	4.6	0.080	0.230*	
[391]		(0.078)	(0.109)	
All Cohorts		0.529**	0.495**	
[391]		(0.096)	(0.105)	
Cohort		0.000	-0.009	
		(0.018)	(0.018)	
Cohort Interaction		-0.091**	-0.049	0.042**
		(0.027)	(0.033)	(0.017)

The cohort label refers to the years in which the father's and the son's earnings began being observed (i.e. '68-'85 Cohort refers to the fathers observed between 1968 and 1972 with sons observed between 1985 and 1989). Heteroskedasticity-robust standard errors are in parentheses in the middle two columns. Bootstrapped standard errors are in parentheses on the difference term in the last column. + significant at 10%; * significant at 5%; ** significant at 1%. The number in the square brackets is the sample size. Earnings are age-adjusted.