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**Moving Out: Transition to Nonresidence among Resident Fathers
in the U.S., 1968 - 1997***

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ABSTRACT

This paper provides the first individual-level estimates of the change over time in the probability of nonresidence for initially resident fathers in the U.S. Drawing on the 1968 - 1997 waves of the Panel Study of Income Dynamics, we use discrete-time event history models to compute the probabilities of nonresidence for six 5-year periods. Our sample consists of men ($N = 1,388$) who are coresident with their biological children at the time of birth. We find that the observed probability of nonresidence doubled over the three decades of the study period, but not linearly. The risk increased substantially in the 1980s and then stabilized in the 1990s. Our multivariate models show that the stabilization was due to changes over time in characteristics such as income; had these remained constant, the likelihood of nonresidence would have increased throughout the study period. Both fathers' and mothers' incomes reduce the likelihood of paternal nonresidence, as do mothers' employment hours.

One of the most visible changes in family life during the last three decades has been the increase in nonresident fatherhood and the accompanying rise in single-parent, predominantly mother-only, households. This trend has aroused substantial scholarly interest and public concern because, on average, children fare better when they live with both biological parents (see Marsiglio, Amato, & Lamb, 2000, for a review of findings). Yet we do not have an answer to the basic question: What is the probability that fathers who initially live with their children will experience nonresidence, and how has that probability changed over time? To date there has been no study of the trend in fathers' transitions from living with their biological children to living apart from those children. This gap is particularly striking because most children are living with their biological fathers at the time of birth (Bumpass & Lu, 2000). Thus the modal experience of nonresidence for fathers is the transition from residence with their biological children to nonresidence.

Although the typical transition to nonresidence for initially resident fathers occurs in the context of union disruption, it is difficult to determine the probabilities of the transition to nonresidence directly from the disruption literature for two reasons. First, not all men in unions are fathers, as one-quarter of married couples during the prime childbearing ages of 20 to 44 do not have children present (Fields & Casper, 2002). Second, recent studies of union disruption do not account explicitly for the presence or absence of children and therefore for men's fatherhood status. Prior work using data from the 1980s indicates that dissolution rates are lower among parents (Heaton, 1990; Lillard & Waite, 1993; Tzeng, 1992), suggesting that it is important to distinguish unions with and without children. Although recent studies show that divorce rates in the U.S. have leveled off (Goldstein, 1999; Raley & Bumpass, 2003; Schoen & Standish, 2001) or have investigated changes in the individual-level risk factors for marital disruption (Heaton, 2002; South, 2001; Teachman, 2002), none identifies the influence of biological fatherhood on the risk of marital disruption. Finally, the rise in father custody in recent years, though not yet a pronounced development, makes it desirable to estimate the likelihood of father nonresidence directly rather than from analyses of union disruption.

We use the Panel Study of Income Dynamics (PSID) to determine the likelihood of nonresidence for fathers who are living with their children at the time of birth over the period 1968 to 1997. Our study is the first to document this trend for fathers in the U.S. who become nonresident after an initial period of coresidence with their biological children. In addition we establish demographic profiles of fathers who stay with their children and of those who do not.

BACKGROUND

There is a large literature on the effect of the absence of fathers on their children's health, educational, and emotional outcomes (e.g., Amato, 2000; McLanahan, 2001; Wu & Thomson, 2001). There are also cross-sectional descriptions of the sociodemographic characteristics of nonresident fathers compared to resident fathers (e.g., Clarke, Cooksey, & Verropoulou, 1998; Garfinkel, McLanahan, & Hanson, 1998; Sorensen, 1997), and studies of changes over time in the proportions of men living with children (Eggebeen, 2002; Hogan & Goldscheider, 2000). What is missing is an analysis of the proportion and characteristics of fathers in the U.S. who are likely to experience nonresident fatherhood (see Juby & Le Bourdais, 1998, for such an analysis of Canadian fathers). Given the scholarly and policy focus on father absence, it is vital that we have this basic empirical information on changes in individual fathers' lives.

We determine the probability that individual fathers will become nonresident with their biological children after an initial period of residence with those children. We limit our study to initially resident fathers for two reasons. First, the vast majority (83%) of fathers live with their biological children in marital or cohabiting unions for some period of time following birth (Bumpass & Lu, 2000). Marital childbearing accounts for at least two-thirds of births in the U.S. and 40% to 50% of nonmarital childbearing occurs within coresidential unions (Bumpass & Lu, 2000; McLanahan, Garfinkel, Reichman, & Teitler, 2001; Ventura & Bachrach, 2000). Second, there is evidence that what matters for child well-being may be more a matter of instability and change in family structure than father absence per se (Wu, 1996; Wu & Thomson, 2001). Our focus on nonresidence among initially resident fathers also offers advantages with respect to data quality, which we discuss below.

In contrast to prior static, cross-sectional comparisons of nonresident and resident fathers, our analysis follows a sample of men from the start of their coresidential relationships with their children to the termination of that coresidence. The longitudinal nature of the PSID data allows us to determine the probability of transition into nonresidence for initially resident fathers. We also examine the change in that probability over the last three decades and provide a demographic profile of the kinds of fathers most likely to experience this event.

Our sample consists of 1,388 men who became fathers throughout the period 1968 to 1997. We include fathers who were resident with their biological children when those children were born, and model the probability of their becoming nonresident fathers subsequently. Our main independent variables are groups of years that capture the period changes in that probability; others are those conventionally included in analyses of union disruption, such as union duration, race, religion, and income. In contrast to the most recent studies of marital disruption, we include number of children and age of youngest child to determine whether children's characteristics affect the risk of nonresidence.

The sample includes about 100 fathers who were cohabiting when they had their first biological children. Childbearing is increasingly common in cohabiting unions (Bumpass & Lu, 2000), and ideally we would compute probabilities of nonresidence separately for cohabiting fathers. But because the PSID does not identify them explicitly before 1983, and because their number is small, we do not report results separately for cohabiting fathers. In tests that we do not document here, we find that our analysis is unaffected by the exclusion of cohabiting fathers because they are a small proportion of the sample. We note however that cohabiting fathers after 1983 are more likely to experience nonresidence than married ones (results not shown).

Our focus on initially resident fathers mitigates some of the problems of research on nonresident fathers using available survey data. Substantial proportions of nonresident fathers are underrepresented in surveys due to either sampling frame undercoverage or survey nonresponse (Rendall, Clarke, Peters, Ranjit, & Verropoulou, 1999). For example, African American nonresident fathers are underrepresented in national surveys such as the National Survey of Families and Households (NSFH) and the Survey of Income and Program Participation (SIPP) (Clarke et al., 1998; Garfinkel et al., 1998; Sorensen, 1997). A second problem is that, even when participating in surveys, men appear to underreport absent children (Bachu, 1996; Clarke et al., 1998; Seltzer & Brandreth, 1994). Sorensen (1997) and Garfinkel et al. (1998) estimate that the NSFH suffers a 40% overall deficit of nonresident fathers, and Sorensen shows that about two-thirds of that deficit is due to the underreporting of absent children. A third difficulty is that some surveys do not allow for direct identification of nonresident fathers or do not gather data on nonresident fathers (e.g., Current Population Surveys). The SIPP, for

example, requires indirect identification of nonresident fathers based on questions about financial payments to children living elsewhere, fertility, and household composition (Sorensen 1997).

A recent study by Rendall et al. (1999), however, suggests that the severity of these problems is reduced substantially by using panel data and explicitly focusing on men who have lived with the mothers of their children. Using the PSID, the authors compare retrospective fertility data with yearly information available from the panel aspect of the survey. They find that men's reporting deficits for births outside of marriage are of substantial magnitude, ranging from one-third to one-half. They find that if one begins with cohabitational fertility (i.e., children born while the father and mother are living together), however, one can quite accurately identify nonresident fathers. Further, tracking initially resident fathers diminishes problems of nonresponse and noncoverage of absent fathers, and permits a more accurate estimation of the socioeconomic characteristics of nonresident fathers (i.e., income) because those variables can be measured before the possible attrition of fathers following nonresidence.

METHOD

Data

Starting from Rendall and colleagues' recommendation, we follow men from the time they become resident fathers until they experience nonresidence or are censored. For reasons discussed below we cannot eliminate completely the use of retrospective fertility data, but our focus on initially resident fathers reduces the severity of problems associated with its use. Our fertility and household residence data come from the 1968 - 1997 PSID. The PSID began in 1968 with 18,000 individuals in 5,000 households. Those individuals have been interviewed every year since then, along with their current coresidents (i.e., partners, children, and others) even after they left their 1968 households. The first wave oversampled low-income households, so that there is a large subsample of African Americans; over a quarter of our sample consists of African Americans. In 1990 the PSID added a sample of 2,000 Latino households; unfortunately our analyses omit this sample because it has not been available across the entire survey period. We use the early release individual dataset that contains the final version of PSID data for the years 1968 to 1992, and a preliminary version of the data for the years 1993 to 1997.

The PSID has household composition data for the fathers in our sample for every year from 1968 to 1997, allowing us to determine whether fathers and children are resident in the same household in a given year. Ideally we would also have annually updated fertility data for the entire period so we could establish coresidential biological fatherhood without ambiguity. The PSID did not collect fertility data before 1985, however. Thus, prior to 1985, we cannot use the main PSID dataset to identify unambiguously the biological children of fathers. Instead, we draw on the Childbirth and Adoption History supplement, added in 1985. This file is the most complete accounting of individuals' birth and adoption histories available to PSID users, containing retrospective fertility data from 1968 to 1985 and updated data annually following 1985. (See the PSID website, <http://www.isr.umich.edu/src/psid>, for descriptions of the supplement as well as the main PSID data.)

The partially retrospective character of the Childbirth and Adoption History supplement means that we cannot eliminate entirely nonresponse and noncoverage problems. We draw on it because, unlike the main dataset, it unambiguously identifies biological parent-child pairs for the years 1968 to

1984. The supplement excludes men not present for the retrospective interview in 1985, however. If these men are disproportionately likely to experience nonresidence, we may underestimate the probability of nonresidence in earlier years. The issue of fathers underreporting fertility outside marriage (or cohabitation) in retrospective data does not affect our analysis, however. Crucially, too, the panel-updated household composition data allow us to determine fathers' incomes, employment, and other important covariates *before* fathers experience nonresidence. On balance, therefore, our procedure reduces substantially the problems associated with cross-sectional or purely retrospective data.

We combine the retrospective fertility data before 1985 and the annually updated fertility data after 1985 with the panel-updated household composition data for the entire period. In this way the Childbirth and Adoption History supplement, when linked with the main PSID individual and family data, allows us to follow the same men over three decades, as they have children, live with them, and are residentially separated from them. We also obtain information on the characteristics of fathers, mothers, and children from the main PSID dataset for the entire period 1968 to 1997. Our data allow us to determine (a) the likelihood that coresident, biological fathers become nonresident; and (b) the change in this likelihood from 1968 to 1997. We also use sociodemographic variables such as race, education, income, and age to identify the characteristics of men who are more or less likely to experience nonresidence with their biological children.

Sample

Our sample consists of all male household heads in the PSID identified as being biological fathers in the child history data supplement. New men entered the PSID in every year of our study period, and men present in earlier years became fathers in subsequent years. As a result, new fathers are added to our sample in every year from 1968 to 1996. Given that our focus is on nonresidence in the context of father-child relationships that start out as coresidential ones, we include only fathers who are living with their children as well as the children's mothers when the children are born. We restrict the sample to household heads because they are more likely to stay in the survey than nonheads, and because they have the most complete information on variables such as income and employment hours. (Our results are not very different when we include about 500 nonhead fathers in the sample. For details, see the section "How robust are the findings?") We exclude men with first births before 1968 because we do not have their residential histories before that year. It is possible that these men experience nonresidence with their children before we observe them for the first time in 1968. We also exclude about a dozen men who have their first child in 1997, the last year of our study period.

Our selection criteria leave us with a sample of 1,388 men, of whom 26% are African American; the rest are White. As already noted, we exclude members of the Latino sample initiated in 1990, because these sample members are not present throughout the survey period. We also exclude 34 men who are members of other racial and ethnic groups, such as Native Americans and Asian Americans, because their small numbers do not allow us to draw representative conclusions.

Our units of analysis are person-years of coresidence. Each father contributes a person-year, or spell, of coresidence for each year he lives with his children. We define a spell of coresidence as a one-year period at the beginning of which the father is known to be living with his child or children. A father begins contributing spells as soon as he has his first biological child in the context of a coresidential union. This can happen in any year between 1968 and 1996, so that new fathers enter the sample in

every year during this period. Once a father starts contributing spells to the risk set, he continues to do so as long as he is coresident with any biological child from the same coresidential union who is less than 18 years old. Thus, he can contribute spells for the entire period of study if he has multiple children with the same mother. He ceases to add spells when he experiences nonresidence with a child less than 18 years of age. We focus on the probability of the first instance of nonresidence; that is, a father stops contributing spells after his first experience of nonresidence even if he becomes a resident father subsequently, whether with the same children or with children he has in the context of a new union. If he does not become nonresident, he stops contributing spells in 1997 or when his youngest child turns 18, whichever comes first. He also stops adding spells if and when he drops out of the survey or dies. Nineteen percent of the fathers in our sample are lost to attrition, and about 2% to mortality. The latter figure is an upper bound and includes fathers who die after their youngest children have turned 18. Because we are concerned only with voluntary transitions to nonresidence, and because the proportion of fathers who die before their children turn 18 is small, we do not model separately the likelihood of nonresidence through death. The total number of spells contributed by all fathers in the sample is 14,320.

Variables

Our dependent variable is a dichotomous measure for nonresidence that equals 1 in the person-year a given father experiences nonresidence, and equals 0 in all other person-years. Our main independent variables are indicators for each of the 5-year periods, with the period 1968 - 1972 serving as the reference category. In addition we control for the variables typically used in the union disruption literature, such as age, race, income, religion, and union duration (see White, 1990 for a review). Given that younger men are more likely to experience union disruption, we include the father's age when he has his first child as a continuous variable (models using categorical age variables yield the same results). We use this measure rather than age at union formation because it is more directly related to fatherhood. Using age at union formation instead does not change our findings. For reasons noted above, our models are restricted to Whites and African Americans, and we include the father's race as a dichotomous variable identifying these two groups. Several studies of marital disruption explicitly determine the effect of being Catholic. Accordingly we add an indicator that distinguishes Catholics from non-Catholic Christians, and from members of other religions, of whom there are very few in our sample. We also control for union duration because it is known that the likelihood of marital disruption decreases with marital duration. We measure duration from the year of marriage for fathers who are married when they enter the sample, and from the year of first observed coresidence with their partners if they are cohabiting.

We also control for fathers' and partners' income, employment hours, and education; these variables are lagged by one year. Our inclusion of income is motivated by the large literature on the effect of both men's and women's incomes on union stability, though their effects have not been unambiguously established (e.g., Brines & Joyner, 1999; Dechter, 1992; Ono, 1998). We control for both father's and partner's total earnings in constant 1983 dollars. We also add both father's and partner's employment hours in the previous year in the form of continuous variables. In addition we control for both father's and partner's education. Because the PSID's education measures are not exactly comparable across all years, we collapse years of education into dichotomous variables

indicating whether they have had any years of college. Finally, we include certain characteristics of the fathers' children. We control for the age of the father's youngest child to determine whether fathers are less likely to become nonresident with younger children than older ones. It is also possible that the probability of nonresidence decreases with the total number of children, and we include that number as a control.

Analytic strategy

We use discrete-time event history analysis to estimate the probability of nonresidence. (Separate continuous-time analyses give us essentially the same results.) This amounts to performing a logistic regression in which the units of analysis are person-years of coresidence, and the dependent variable is the dichotomous measure of father's residential status discussed above (Allison, 1984). Our model is:

$$\log(P(t, \mathbf{X}) / (1 - (P(t, \mathbf{X}))) = \sum \mathbf{a}_j(t) + \sum \mathbf{b}_k X_k + \sum \mathbf{b}_l X_l(t)$$

Here $P(t, \mathbf{X})$ is the probability of nonresidence at time t , given the values of the independent variables \mathbf{X} . The period indicators are represented by the \mathbf{a}_j 's. The X_k 's are the time-independent covariates in the model such as father's age at the time he enters the sample, and the $X_l(t)$'s are the time-dependent covariates such as income. The values of fixed covariates, such as age at the time of entry into the risk set, are the same for all spells contributed by a father. For time-varying covariates such as income, we attach the values in a given year to the spells that begin in that year. Following other recent analyses using the PSID, we perform unweighted analyses because the sample selection probability weights for each individual change across the years (Brines & Joyner, 1999; Ono, 1998; South, 2001).

The coefficients from the model give us the changes in the annual log odds of nonresidence for unit increases in the independent variables. We determine the period trend in nonresidence by estimating the effects on the log odds of nonresidence for six consecutive 5-year periods: 1968 - 72, 1973 - 77, 1978 - 82, 1983 - 87, 1988 - 92, and 1993 - 96. Note that the last period consists of four rather than 5 years, because we leave out the men who become fathers in 1997. We also report the observed proportions of fathers who experience nonresidence in each period.

RESULTS

Descriptive data

Table 1 shows, for each period, the number of resident fathers, the number of person-years they contributed, and the observed proportion of fathers who experienced nonresidence. The proportion of resident fathers who made the transition to nonresidence grew from 5 % in 1968 - 72 to 13% in 1983 - 87, and then declined to 10% by 1993 - 96. Note that the resident fathers in each period include fathers who survived the previous period (i.e., did not experience nonresidence during the previous period) as well as men who became biological fathers for the first time in a given period, shown separately. In each period, fathers from earlier periods outnumbered new fathers added in that period. Table 1 also displays the fathers' characteristics and other independent variables measured at the beginning of each period. We see that fathers' mean age, measured at the start of each period, grew

from 25 years to 35 years over the three decades of our study. This is due both to the survival of fathers from earlier periods and because the mean age of new fathers increased by 4 years, from 25 to 29 years. (Though Table 1 shows fathers' mean age in each period, our analyses use fathers' age at the time they enter the sample. This makes no difference to our results.)

Period effects on the probability of nonresidence

Table 2 shows the results of our multivariate analyses. We use three models, all of which employ indicator variables for 5-year groups of years marking the start of spells. The reference period is the first, which includes spells that began in 1968 through 1972. Model 1 includes only these period indicators. Model 2 adds fathers' characteristics, both constant ones such as race and time-varying ones such as income. Model 3 adds their partners' characteristics and the duration of their unions. We report both the coefficients from the models as well as the corresponding exponentiated coefficients, or odds ratios. The units of analysis for all three models are person-years of residential experience. Therefore the coefficients in all models represent additive effects on the annual log of the odds of nonresidence.

The pattern of period coefficients and odds ratios in Model 1 is the same as the observed trend in nonresidence shown in Table 1. That is, they fall slightly in 1973 – 77, then rise until the period 1983 – 87, and fall after that. Separate tests show that the coefficients for the last three periods, 1983 – 87 through 1993 – 96, are significantly different at $p < 0.01$ from the coefficients of the early periods 1973 – 77 and 1978 – 82. The coefficients for the last three periods are not significantly different from one another, however. This means that in the absence of controls for father, mother, and child characteristics, the odds of nonresidence increase until 1983 – 87 with a small dip in 1973 – 77 and then stabilize or decrease in the period 1988 – 97.

Adding the fathers' characteristics in Model 2 alters the period pattern of coefficients. In this model, the coefficients increase in every period following 1973 – 77, though in successively smaller increments. Separate tests show that, as in Model 1, the coefficients of the last three periods are significantly different at $p < 0.01$ from the coefficients of the periods 1973 – 77 and 1978 – 82, but that the coefficients for the last three periods are not significantly different from one another. Model 3 adds partners' characteristics and union duration. In this final model also the coefficients increase in every period following the second, and the increases get successively smaller. Again the coefficients of the last three periods are significantly different at $p < 0.01$ from the coefficients of the periods 1973 – 77 and 1978 – 82. The coefficient of the last period is not significantly different from that of the immediately preceding period 1988 – 92, though it is different from that of the period 1983 – 87 at $p < 0.05$.

We can summarize the results of Models 1 through 3 regarding the period trend in the odds of nonresidence as follows. If we do not take into account the changes over time in the covariates, as in Model 1, the trend in the period coefficients parallels the pattern of observed proportions nonresident shown in Table I. The annual odds of nonresidence increase through 1983 – 87 and decline after that. Upon the addition of the covariates in Models 2 and 3, the odds continue to increase through the final period 1993 – 97. The coefficient for that last period, however, is not significantly different from the coefficient of the penultimate period 1988 – 92.

TABLE 1. DESCRIPTIVE STATISTICS, BY FIRST YEAR OF PERIOD (N = 1,388 fathers)

| | 1968 | 1973 | 1978 | 1983 | 1988 | 1993 |
|--|-------|-------|-------|-------|-------|-------|
| <i>Sample</i> | | | | | | |
| Total number of resident fathers | 164 | 377 | 608 | 872 | 943 | 864 |
| Number of new resident fathers | 164 | 222 | 255 | 315 | 252 | 180 |
| Person-years | 451 | 1368 | 2432 | 3441 | 3734 | 2894 |
| Proportion experiencing nonresidence | 0.049 | 0.042 | 0.076 | 0.130 | 0.122 | 0.102 |
| Average no. of person-yrs per father | 2.75 | 3.63 | 4.00 | 3.95 | 3.96 | 3.35 |
| <i>Father's characteristics at start of period</i> | | | | | | |
| Age, years | | | | | | |
| <i>M</i> | 24.9 | 26.1 | 28.4 | 30.6 | 33.1 | 35.3 |
| <i>SD</i> | 4.2 | 4.5 | 4.9 | 5.8 | 6.4 | 7.1 |
| minimum | 17.0 | 18.0 | 18.0 | 19.0 | 19.0 | 16.0 |
| maximum | 41.0 | 45.0 | 50.0 | 59.0 | 61.0 | 60.0 |
| White: 0 = no , 1 = yes | 0.87 | 0.81 | 0.78 | 0.75 | 0.77 | 0.79 |
| Catholic: 0 = no , 1 = yes | 0.19 | 0.21 | 0.25 | 0.26 | 0.27 | 0.28 |
| College education: 0 = no , 1 = yes | 0.44 | 0.45 | 0.51 | 0.51 | 0.58 | 0.63 |
| Labor income, thousands (1983 dollars) | | | | | | |
| <i>M</i> | 19.0 | 21.0 | 22.3 | 20.5 | 23.8 | 29.0 |
| <i>SD</i> | 9.0 | 13.9 | 13.2 | 14.1 | 20.2 | 32.0 |
| Employment hours | | | | | | |
| <i>M</i> | 44.9 | 44.7 | 43.7 | 42.6 | 44.2 | 44.4 |
| <i>SD</i> | 12.2 | 9.3 | 11.3 | 13.0 | 12.5 | 11.9 |
| <i>Mother's characteristics at start of period</i> | | | | | | |
| College education: 0 = no , 1 = yes | 0.21 | 0.34 | 0.41 | 0.45 | 0.56 | 0.61 |
| Labor income, thousands (1983 dollars) | | | | | | |
| <i>M</i> | 4.1 | 3.8 | 4.7 | 6.0 | 8.0 | 11.6 |
| <i>SD</i> | 5.4 | 5.8 | 6.3 | 7.3 | 8.7 | 17.3 |
| Employment hours | | | | | | |
| <i>M</i> | 22.5 | 19.9 | 20.6 | 20.0 | 25.5 | 26.5 |
| <i>SD</i> | 19.0 | 18.6 | 18.6 | 18.6 | 18.0 | 17.8 |
| <i>Child characteristics at start of period</i> | | | | | | |
| Age of youngest child, years | | | | | | |
| <i>M</i> | 0.0 | 0.8 | 1.6 | 2.6 | 3.8 | 4.9 |
| <i>SD</i> | 0.0 | 1.3 | 2.3 | 3.4 | 4.3 | 4.8 |
| Number of children | | | | | | |
| <i>M</i> | 1.0 | 1.3 | 1.6 | 1.7 | 1.8 | 1.8 |
| <i>SD</i> | 0.2 | 0.5 | 0.8 | 0.9 | 0.9 | 0.9 |
| <i>Union duration, years</i> | | | | | | |
| <i>M</i> | 2.3 | 3.7 | 5.3 | 6.9 | 8.7 | 10.3 |
| <i>SD</i> | 2.2 | 2.7 | 3.9 | 5.2 | 6.4 | 7.2 |

TABLE 2. LOGISTIC REGRESSION MODELS FOR THE RISK OF NONRESIDENCE
(*N* = 14,320 person-years)

| | Model 1 | | Model 2 | | Model 3 | |
|--|---------|----------------|---------|----------------|---------|----------------|
| | b | e ^b | b | e ^b | b | e ^b |
| <i>Is a father in:</i> | | | | | | |
| 1968 - 72 (reference category) | | | | | | |
| 1973 - 77 | -0.423 | 0.655 | -0.446 | 0.640 | -0.406 | 0.666 |
| 1978 - 82 | 0.065 | 1.068 | 0.059 | 1.061 | 0.262 | 1.299 |
| 1983 - 87 | 0.631 | 1.880 | 0.610 | 1.840 | 0.932 | 2.540 ** |
| 1988 - 92 | 0.565 | 1.760 | 0.630 | 1.878 | 1.193 | 3.298 ** |
| 1993 - 97 | 0.552 | 1.737 | 0.686 | 1.986 | 1.311 | 3.709 *** |
| <i>Age at entry into sample</i> ^a | | | -0.046 | 0.955 *** | -0.030 | 0.970 * |
| <i>Father's characteristics</i> ^b | | | | | | |
| White | | | -0.664 | 0.515 *** | -0.732 | 0.481 *** |
| Catholic | | | 0.109 | 1.115 | 0.072 | 1.075 |
| College education | | | -0.047 | 0.954 | 0.061 | 1.063 |
| Labor income (\$1,000s) | | | -0.024 | 0.976 *** | -0.020 | 0.980 *** |
| Employment hours | | | -0.001 | 0.999 | 0.003 | 1.003 |
| <i>Mother's characteristics</i> | | | | | | |
| College education | | | | | 0.187 | 1.205 |
| Labor income (\$1,000s) | | | | | -0.096 | 0.908 *** |
| Employment hours | | | | | -0.073 | 0.930 *** |
| <i>Children</i> | | | | | | |
| Age of youngest child | | | | | 0.137 | 1.147 *** |
| Number of children | | | | | -0.058 | 0.944 |
| <i>Union duration (years)</i> | | | | | | |
| Constant | -4.014 | 0.018 *** | -1.853 | 0.157 *** | -1.470 | 0.230 ** |
| Log likelihood | -1760.0 | | -1699.3 | | -1373.2 | |
| χ^2 | 31.2 | | 152.6 | | 804.7 | |
| <i>DF</i> | 5 | | 11 | | 17 | |

e^b is exponentiated b.

^a Father's age at time of birth of first biological child, in years. ^b White, Catholic, and College education are coded 1 for yes, 0 for no.

* *p* < 0.05; ** *p* < 0.01; *** *p* < 0.001.

We conclude that the underlying propensity of nonresidence increased over the duration of our study, net of the covariates, though the rate of increase slowed in the later periods. At the same time the observed proportions of nonresident fathers decreased in the second half of the study period. This discrepancy between the observed period pattern and the one evident in Models 2 and 3 can be explained by considering the changes over time in the covariates in those models. Fathers' characteristics and those of their spouses and partners changed substantially over the three decades of our study. For example fathers' mean earnings increased from \$19,000 to \$29,000, and their partners' earnings tripled from \$4,000 to \$12,000. Had these characteristics remained constant, the proportions of resident fathers experiencing nonresidence would have increased throughout the study period, though at a slower rate after 1983 – 87. To confirm this we computed the predicted annual rates of nonresidence from Model 3 holding the covariates constant at their 1968 levels (results not shown). Those rates increase throughout the period 1968 to 1997, with the exception of a small decline in 1973 – 77.

Next we turn our attention to the effects of the covariates in the full model, Model 3. We note that men who become resident fathers at older ages are less likely to experience nonresidence. For small increases in age, the annual odds of nonresidence decrease by 3% each year; the odds for men who are 30 when they have their first child are 26% lower than for men who are 20.

Of the fathers' socioeconomic characteristics, race and earnings have powerful effects on the odds of nonresidence, but religion does not matter. The annual odds of nonresidence for White fathers are more than 50% smaller than for African American fathers. And they decrease by 2% for each additional \$1,000 of annual earnings, so that the odds for fathers who earn \$30,000 are about 20% lower than for fathers who make \$20,000. Fathers' employment hours do not affect the odds, nor does their having some college education. In separate models we confirm that the collinearity among father's education, earnings, and employment hours does not affect the magnitudes and significance levels of their respective coefficients.

Certain characteristics of mothers are also associated with father nonresidence: The more mothers earn and the more time they spend on paid work, the lower those odds. An additional \$1,000 of mothers' annual earnings reduces the odds by about 9%, a larger effect than the 2% resulting from the same increment in fathers' earnings. In separate tests we used the ratio of mother's to father's earnings rather than absolute earnings. Our results were the same: Mothers' earnings, whether relative or absolute, reduce the odds of nonresidence for fathers. Further, unlike fathers' employment hours, those of mothers have a dampening effect on the odds, with every additional hour reducing it by 7%. This means that the annual odds for fathers whose partners work 40 hours a week are almost 50% lower than for fathers whose partners do not work outside the home.

We obtain the same results with alternative specifications for income and employment that use categorical rather than continuous variables; we retain the continuous variables for parsimony. Finally, we note that there is some collinearity between mothers' earnings and employment hours. The addition of mothers' employment hours reduces the effect of mothers' earnings on the odds from about 30% for every \$1,000 to the 9% we report in Table 2. We include both variables because each of them has a substantial and significant effect on fathers' odds of nonresidence.

Of the child variables, the number of children does not affect the odds of nonresidence, but age of the youngest child does matter, with every additional year increasing the odds by about 15%. The odds that a father whose youngest child is 10 years old, for example, are almost twice as high as the odds if the youngest child is 5. As expected, union duration decreases the risk of nonresidence, with every additional year of marriage reducing the odds by almost 10%. There is substantial collinearity between union duration and age of youngest child; however, separate tests show that this does not affect the period coefficients or the predicted probabilities.

How robust are the findings?

The generalizability of our results is subject to a few caveats. The most important of these is sample attrition over time and its possible correlation with nonresidence. Like all longitudinal surveys, the PSID loses respondents to follow up every year. The fathers who drop out from the survey may be more likely to experience nonresidence than the fathers who remain. Indeed, some of them may leave the survey precisely when they become nonresident. In this case the true probabilities of nonresidence would be higher than our estimates suggest. This is especially a concern for the later periods: Higher probabilities in those periods could mean that the risk of nonresidence has not stabilized in the 1990s.

To assess how attrition affects our results, we re-estimate our models using different assumptions regarding the rates of nonresidence among fathers who are lost to the survey. First, we suppose that the fathers who are lost to attrition in a particular period experience nonresidence at the same rate as all fathers in that period. We randomly designate that proportion of attrited fathers as nonresident; that is, we assume that they experience nonresidence rather than attrition in the last year that they are in the survey. In the last period, for example, 3% of all fathers experience nonresidence. Accordingly we randomly code 3% of the attrited fathers as nonresident, and re-estimate our models including these fathers. We find that the coefficients in our model change only slightly, and their time trend is preserved. Next we assume that the fathers who drop out of the survey experience nonresidence at three times the rate as fathers who remain. Again our coefficients change very little, and their pattern over time is unaffected. Our findings are therefore robust to fairly conservative assumptions about the rates of nonresidence among attrited fathers.

Second, we consider the implications of limiting our sample to PSID household heads. We exclude 510 nonheads because we lack information on their incomes, employment hours, and other covariates. We do have their fertility and residence histories, however. Because it is possible that fathers who are not household heads experience nonresidence at different rates from fathers who are, we add the nonheads and re-estimate Model 1 in Table 2. Our period coefficients differ slightly from the ones that appear in Table 2, but the period trend is unaffected. The exclusion of nonheads from our sample, therefore, does not affect our findings regarding the period trend in the probability of nonresidence.

CONCLUSION

We used longitudinal fertility and household membership data from the PSID to document the trend in nonresidence for men who initially lived with their biological children. Our focus here is not on union dissolution per se, but on the resulting separation of men from their biological children, because that has been identified as one of the most troubling consequences of union dissolution (e.g., McLanahan & Sandefur, 1994). There is accumulating evidence that instability, in itself, has negative outcomes for children (Wu, 1996; Wu & Thomson, 2001).

We find that biological resident fathers were twice as likely to experience nonresidence during the second half of the 1968 - 97 period than during the first. The increase in observed likelihood of nonresidence was not linear, however. It occurred in three stages: (a) little change for most of the 1970s, (b) a sustained increase starting in the late 1970s and continuing through the late 1980s, and (c) a leveling off during the 1990s. Our results are broadly consistent with other recent research showing an increase in the likelihood of marital disruption in the 1970s and 1980s, followed by a steadying in the 1990s (Casper & Bianchi, 2001). Our multivariate models show that controlling for changes in father and partner characteristics, however, the annual odds of nonresidence increased throughout the duration of our study, with the rate of increase slowing in the latter periods. Therefore, had those characteristics remained constant, the proportions of initially resident fathers experiencing nonresidence would have increased throughout the three decades. The limitations of our analysis include our reliance on retrospective fertility data for part of the study period, and sample attrition over that period.

There are several implications of our findings. First, we think they are grounds for both concern and cautious optimism. Most generally, our findings suggest that even brief historical periods can contain considerable fluctuation. On one hand, the probability of nonresidence for fathers has increased substantially over the last three decades, consistent with Teachman's characterization of the "pervasive impact of historical period" on the risk of marital disruption (Teachman, 2002, p. 346). On the other, this increase has not occurred in the implacable fashion that suggests inevitability.

Second, our study adds to a large and growing literature linking family instability to economic hardship. We find that the probability of nonresidence decreases significantly as both fathers' and mothers' incomes increase. Policymakers concerned with the formation and maintenance of two-parent families, especially when children are involved, should continue to recognize that it is far more likely that fathers remain with their children when family incomes are higher. Our findings regarding women's income and employment add to the growing body of evidence that women's incomes may have stabilizing rather than disruptive effects on family life. Indeed, the rise in women's earnings over the study period contributed to the decline in the observed probability of nonresidence after 1983 - 87.

Finally, our results contribute to our understanding of the intersection of gender and contemporary family patterns in the United States. Several scholars have argued that the current parenting system is one in which men are "serial parents" or are "swapping families" (Furstenberg & Cherlin, 1991; Manning & Smock, 2000). By contrast, women comprise the vast majority of single parents and overwhelmingly live with all of their biological children. Levels of union disruption and nonmarital childbearing have led to increasing numbers of fathers living separately from their biological children, even if the trend has receded somewhat in recent years. At the same time, men frequently go on to remarry and to have new biological children or stepchildren. Manning, Stewart, and Smock (2003) find that roughly half of all nonresident fathers have parenting responsibilities beyond a single set of nonresident children, and that nearly three-quarters of those who are remarried or cohabiting have responsibilities for other children. Thus, even as men become nonresident fathers of some of their biological children, they are forging new ties to other biological and stepchildren with whom they live.

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