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THE DYNAMICS OF NONRESIDENTIAL FATHERHOOD IN THE US, 1968-1997*

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THE DYNAMICS OF NONRESIDENTIAL FATHERHOOD IN THE US, 1968-1997

ABSTRACT

This paper provides the first individual-level estimates of the change in the hazard of nonresidence for fathers in the US. Drawing on the 1968-1997 waves of the Panel Study of Income Dynamics (PSID), we use Cox regression models to compute the relative hazards of nonresidence for six 5-year periods. Our sample consists of men who are coresident with their biological children, and the children's mothers, at the time of birth. We find that the hazard of nonresidence has more than doubled, with the increase concentrated in the 1980s. We also find that the hazard is substantially higher for African-American than for white fathers.

THE DYNAMICS OF NONRESIDENTIAL FATHERHOOD IN THE US, 1968-1997

INTRODUCTION

The structure and practice of parenting have been altered dramatically by changes in family behavior over the past three decades. Increasing rates of nonmarital fertility and marital disruption have eroded the postwar ideal of households consisting of two parents and their biological children. Since women continue to be the primary custodians and caretakers of children, one result of these trends is that a substantial proportion of children do not live with their biological fathers. In 1996, this proportion was approximately one-third (Fields 2001).

This situation has aroused concern in public, policy and scholarly arenas, with fatherhood becoming a prominent social issue. Nicholas Davidson has called the absence of biological fathers in their children's lives "America's greatest social catastrophe," and sees it as the root of crime, drug use and poverty (Davidson 1990; also see Blankenhorn 1996; Popenoe 1996; Whitehead 1996). Even less alarmist commentators suggest that the institution of marriage must be strengthened if we are to avoid the harmful consequences for children of nonresidential fatherhood (Furstenberg and Harris 1992).

The discussion surrounding nonresident fathers, however, suffers from a basic empirical gap, namely a lack of individual-level analyses of trends in the incidence of nonresidential fatherhood. Compared to other family changes, these trends are sparsely documented. Additionally, there are very few individual-level estimates of the likelihood of fathers experiencing nonresidential fatherhood (see Clarke et. al 1998 for an exception).

We address this gap by reporting the first individual-level estimates of changes over time in the hazard of coresidential, biological fathers experiencing nonresidential fatherhood in the US. We also provide a basic demographic profile of the kinds of father most likely to experience this event. While we do not address directly the larger debate on the significance and implications of these trends, our findings provide a necessary empirical baseline for that discussion.

BACKGROUND

Families are increasingly likely to be spread across more than one household, with one biological parent, typically the father, living in a separate household. In 1996 over 20 million children did not live with their biological fathers, and estimates suggest that about half of all children will live without their biological fathers at some point before adulthood (Bumpass and Sweet 1989; Norton and Miller 1992).

In light of these trends, the issue of nonresident fatherhood has attracted much theoretical and empirical attention from social scientists. (See Marsiglio et. al 2000 for a review.) In large part, this attention has been catalyzed by concern for the well-being of children. A growing literature suggests that living apart from biological fathers has negative effects on children, although researchers continue to debate the causal mechanisms accounting for this effect (e.g., Amato 2000; Amato and Booth 1992; Amato and Gilbreth 1999; Amato, Loomis, and Booth 1995; Furstenberg and Kiernan 2001; McLanahan and Sandefur 1994; Seltzer 1994; Wu 1996; Wu and Martinson 1993; Wu and Thomson 2001). The evidence suggests that income accounts for at least some of these effects due to child custody patterns. Since women usually remain the primary caretakers, women's low potential earnings and meager child support mean that nearly a majority of children in households without fathers live in poverty (Garfinkel and McLanahan 1986; Holden and Smock 1991). Accordingly, a national focus on strengthening child support has emerged as one way to improve the well-being of children, with an accompanying interest in nonresident fathers.

The existing demographic research on nonresident fathers has thus been largely focused on the determinants of child support and the consequences of various forms of nonresident father involvement, including economic support, for children's well-being (e.g., Argys et al. 1998; Beller and Graham 1993; Cooksey and Craig 1998; Garfinkel and Robins 1994; Hernandez, Beller, and Graham 1995; Hill 1992; King 1994; Knox 1996; Manning and Smock 1999, 2000; Mott 1990; Paasch and Teachman 1991;

Rangarajan and Gleason 1998; Seltzer 1991, 1994, 1998; Smock and Manning 1997; Sorensen 1997; Stewart 1999). There are also a handful of studies that examine the demographic characteristics of nonresident fathers, but all report difficulties identifying and describing nonresident fathers (Clarke, Cooksey, and Verropoulou 1998; Garfinkel, McLanahan, and Hanson 1998; Seltzer and Brandreth 1994; Sorensen 1997).[1]

The problem is identifying nonresident fathers accurately with available survey data. Substantial proportions of nonresident fathers are underrepresented in surveys due to either sampling frame undercoverage or survey nonresponse (Rendall et al. 1999). Regarding the former, most surveys are restricted to the noninstutionalized population, thus eliminating men in prison, the military, and other kinds of group quarters. In addition, surveys suffer from the same undercounting that the census does, with proportionately larger effects on minority men and those of lower socioeconomic background. 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, McLanahan and Hanson 1998; Sorensen 1997). Survey nonresponse is generally somewhat higher for men than women, and especially those who are not currently married, minority men, and those of lower socioeconomic status.

Another problem is that, even when participating in surveys, men appear to underreport absent children (Bachu 1996; Cherlin, Griffith, and McCarthy 1983; Clarke et al. 1998; Seltzer and Brandreth 1994). For example, both 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. Seltzer and Brandreth (1994) come to a similar conclusion, also using the NSFH.

[1] Clarke et al. (1998) compare nonresident fathers in Britain and the United States. They find that the surveys used for each country are quite similar in undercounting nonresident fathers, suggesting that a comparison is still valid.

A third problem is that some surveys do not allow for direct identification of nonresident fathers or even attempt to gather data on nonresident fathers (e.g., Current Population Surveys). The SIPP, for example, requires several steps simply to identify nonresident fathers, requiring indirect identification based on questions about financial payments to children living elsewhere, fertility, and household composition. Identification of nonresident fathers also requires making some assumptions to eliminate the problem of falsely identifying adult children living elsewhere (Sorensen 1997).

An important recent paper, however, suggests that these problems are minimized by the use of panel data and by an explicit focus on men who have coresided with the mother of their children. Using the Panel Study of Income Dynamics (PSID), Rendall et al. (1999) compared retrospective fertility data with the yearly information available from the panel aspect of the survey. Like other studies, they find considerable problems with the retrospective data. Men's reporting deficits for births outside of marriage, including fertility that occurred in a past marriage or outside of marriage altogether, are of substantial magnitude, ranging from one third to one half. At the same time, they find that if one is interested in cohabitational fertility—fertility occurring while the father and mother are living together—one can quite accurately determine nonresident fatherhood. Rendall and colleagues state:

> With panel collection, men's (marital or nonmarital) cohabitational fertility is recorded within a year of the child's birth, and therefore before union dissolution and possible subsequent nonresponse and noncoverage. Interviewing men while they reside with their children does not, of course, solve the problem of incomplete surveying of fathers in later periods when they are nonresident. However, the estimation of important characteristics of nonresident fathers, such as income, is clearly facilitated by information on their human capital and earnings before their attrition following union dissolution (p. 143).

DATA AND METHODS

Because it is difficult to identify nonresident fathers and obtain accurate information about them, existing research leaves some fundamental empirical questions unanswered. Starting from Rendall and colleagues' recommendation, we draw on retrospective fertility data from 1968 to 1984 and panel-updated data for the years 1985 to 1997 from the PSID to determine (i) the likelihood that coresidential, biological fathers become nonresidential fathers, and (ii) the change in this likelihood from 1968-1997. To date, these probabilities are unknown for the United States (see Peron et al. [1999] for a similar analysis for Canada). 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. These questions are basic, but critical for assessing the significance of recent trends for men's family experiences.

By using panel-updated data for half of our study period, we mitigate considerably the problems associated with the use of retrospective fertility data. We restrict our attention to fathers who are living with the mothers of their children when the latter are born. This focus is both a strength and a limitation. Clearly, the major strength of our approach is that of data quality, as Rendall et al. (1999) underscored. Another strength is that we include fathers in both kinds of coresidential union, cohabitation and marriage, in our analyses. Cohabitation has become commonplace in the United States and is increasingly a site for childbearing; most of the increase in nonmarital childbearing between the 1980s and 1990s stems from births to cohabiting couples (Bumpass and Lu 2000). In particular, cohabitation plays a relatively large role in family formation and fertility among African-Americans. Greater proportions of African-Americans than whites select cohabitation as their first union (Clarkberg 1999; Loomis and Landale 1994; Willis and Michael 1994). Along with Latinos, African-Americans are more likely than whites to give birth to children while cohabiting (Loomis and Landale 1994; Manning forthcoming). Thus, it is important that childbearing within cohabitations is included in any analysis of nonresidential fatherhood.[2]

2 A possible limitation of our focus on fertility in coresidential unions is that we omit the experiences of some African-American men, since African-Americans appear more likely to have children outside of coresidential unions than do members of other racial and ethnic groups (Bumpass and Lu 2000; Raley 2001). On the other hand, the omission of African-American nonresidential fertility is compensated for to some extent by the unusually large proportion of our sample, 27 percent, that is African-American.

Our data come from the 1968 to 1997 waves of the Panel Study of Income Dynamics (PSID), which began in 1968 with 18,000 individuals in 5,000 households. Those individuals have been reinterviewed every year since then, along with their current coresidents—partners, children, and others—even after they leave the 1968 household. The first wave in 1968 oversampled low-income households, so that there is a large subsample of African-Americans. In 1990 the PSID added a sample of 2,000 Latino households. Unfortunately, our analyses omit the Latino sample because it has not been available across the entire survey period. Further, 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.

Our file construction proceeds as follows. First, we identify biological father-child pairs from the Childbirth and Adoption History supplement, which PSID added in 1985. This file is the most complete accounting of individuals' birth and adoption histories available to PSID users. It contains retrospective fertility data on sample members from 1968 to 1985, and annually updated data for subsequent years. (See the PSID website, http://www.isr.umich.edu/src/psid, for descriptions of the supplement as well as the main PSID data.) Second, we obtain information on the fathers, mothers and children from the main PSID dataset. We track these individuals and their trajectories—their histories of coresidence and nonresidence, income, education, socioeconomic characteristics, and so on—over nearly thirty years of PSID data. In this way the supplement, when linked with the PSID individual and family data, allows us to follow the same men over three decades, as they have children, live with them, and are separated from them. A drawback of using the supplement is that we may be underestimating the hazard of nonresidence by excluding the men who were not present for the retrospective interview in 1985, if these men are disproportionately likely to experience nonresidence. Our estimates of the hazard of nonresidence may also be biased downward by the deficit in nonmarital and previously married fertility in the retrospective portion of the supplement (Rendall et al. 1999). This problem is less serious, however, than that of overestimating the hazard. The limitations of the supplement actually make our estimates of the hazards of nonresidence conservative; the true hazards may be higher.

Sample

Our sample consists of all male household heads in the PSID who are identified as being biological fathers in the child history data supplement. Further, they are coresident 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 non-heads, and because they have the most complete information on variables like income and employment hours. 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 leave out the dozen or so men who have their first child in 1997, since they cannot contribute complete spells. Finally, because our focus is on nonresidence in the context of father-child relationships that start out as coresidential ones, we eliminate fathers whose children are not coresident with them at birth. We include fathers in both marital and cohabiting unions and cannot distinguish between them for the entire study period, because the PSID does not explicitly identify cohabiting partners before 1983. Because of the way cohabitors are identified in the PSID, our sample includes only cohabitations that last for more than a year.

Our selection criteria leave us with a sample of 1,512 men, of whom 404, or 27 percent, are African-American; the rest are white. As already noted, we exclude members of the Latino sample initiated in 1990, since 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 about their hazards. Our results for whites and African-Americans are not affected by this exclusion.

Variables

To ascertain the period trends in the hazard of nonresidence, we include dichotomous variables for each of the five-year periods, with the period 1968 to 1972 serving as the reference category. We control for key demographic and socioeconomic characteristics of the fathers and their partners. We include the father's age when he enters risk, i.e. when he has his first child, as a continuous variable.[3] For reasons noted in the sample description above, our models are restricted to whites and African-Americans; accordingly the father's race is included as a dichotomous variable distinguishing these two groups. To determine whether fathers are less likely to become nonresident with younger children than older ones, we control for the age of the father's youngest child. It is also possible that the hazard of nonresidence decreases with the total number of children; thus we include that number as a control.

We also control for fathers' and partners' income, employment hours, and education; these variables are lagged by one year. A large literature on the impact of incomes on union stability suggests two possibilities, described as the 'independence' and 'income effect' hypotheses (e.g. Becker 1981; Dechter 1992; Ono 1998; Oppenheimer 1997; Ruggles 1997). The first suggests that the likelihood of marital dissolution increases with wives' incomes, both absolute and relative to their partners'. The second proposes that wives' incomes stabilize marriages by reducing financial tensions. To account for the possible effects of income on the hazard of nonresidence, we include both father's and partner's total earnings. We also control for both father's and partner's employment hours in the previous year in the form of continuous variables. In addition, we determine whether government income supplements affect the hazard of nonresidence by specifiying whether the household receives Aid to Families with Dependent Children (AFDC) support. Because of the small number of households receiving support in any given year, we control for AFDC support as a categorical variable. Finally, we control for both father's and partner's education. Since the PSID's education measures are not exactly comparable across all thirty years, we collapse years of education into dichotomous variables indicating whether or not they have had any college experience. This also reduces the collinearity between education and earnings in our models.

3 Separate models with categorical age variables do not yield substantively different results.

Analytic Strategy

We use Cox regression models to obtain our estimates of the hazards of nonresidence during the period 1968 to 1997. Our dependent variable is the hazard of nonresidence at time t, which is given by the following expression:

$$h(t, \mathbf{X}) = h_o(t)\exp(\sum b_i X_i + \sum b_j X_j(t))$$

Here $h_o(t)$ is the baseline hazard faced by all fathers at time t, and is not estimated directly. The X_i's are the time–independent covariates in the model, and the $X_j(t)$'s are the time–dependent covariates.

A PSID male enters the risk set 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; therefore, new fathers enter the risk set every year during this period. A father remains at risk as long as he has any coresident child who is less than 18 years old. If he has multiple children, he can remain at risk for the entire period of our study. He exits the risk set as soon as he experiences nonresidence with any child less than 18 years of age. Otherwise, he is treated as censored, and either stays on until the final year, 1997, or is lost to attrition before then. He is also censored when his youngest child turns 18, if he has not experienced nonresidence before then, whether or not he does so subsequently.

Our units of analysis are person-years of coresidence. Each father contributes a person-year, or spell, of coresidence for each year in which he is coresident with his children.[4] He ceases to do so after he experiences nonresidence, or is lost to the survey without experiencing the event. We define a spell of coresidence as a one-year period at the beginning of which the father is known to be coresident with his child or children. To determine the risk factors associated with nonresidence, we attach covariates to the spells. 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, like income, we attach the values in a given

4 Since our focus is on the first instance of nonresidence, we ignore the years of coresidence experienced by a small number of fathers

year to the spells that begins 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 and Joyner 1999, Ono 1998, South 2001).

We analyze the trend in nonresidence by determining the hazards of nonresidence for consecutive 5-year periods of calendar time, defined by the years in which spells begin. This gives us six periods: 1968-72, 1973-77, 1978-82, 1983-87, 1988-92, and 1993-96. (Note that the last period consists of four rather than five years, since we leave out the men who become fathers in 1997.) Our analysis is designed to determine the relative hazards of nonresidence faced by a father in each of these periods; the same father can be at risk in multiple periods. We answer the question: are fathers more likely to experience nonresidence in the years 1988-92, say, than they are in 1978-82? To prevent confounding with age, we control for father's age at the birth of the first child.

We estimate two multivariate models. The first includes just the period indicator variables and a control for fathers' ages at the time they enter risk. We control for their ages at the time they enter risk to account for the increase over the study period in fathers' ages at the birth of their first children, the event that causes them to enter the risk set.[5] The second model adds the child variables (age of youngest child, number of children) and socioeconomic indicators (father's race, both parents' earnings, employment hours, and education).

RESULTS

Descriptives

Table 1 presents summary information on all variables in our models. The descriptives for the time-independent variables refer to fathers, while those for the time-dependent variables refer to person-years of coresidence. Of the 1,512 fathers in our sample, 482, or 32 percent, experience the event of nonresidence

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following the first event of nonresidence.

5 A separate model excluding father's age at risk entry yields the same pattern of period coefficients as Model 1, though the individual coefficients are slightly different.

by 1997. The proportion of whites experiencing nonresidence is 27 percent, while that of African-Americans is 45 percent. Another 19 percent of the sample, made up of almost identical proportions of whites and African-Americans, is lost to the survey without experiencing the event by the year of attrition. The rest continue to be coresident with their children from the time they enter the risk set until 1997, or experience nonresidence only after their youngest child turns 18. The mean time at risk is just under 10 years, and the median number of years at risk is 8.5.

TABLE 1. DESCRIPTIVES

	N	Mean	SD
Experiences nonresidence	1,512	0.32	
White	1,108	0.27	
African-American	404	0.45	
Time-independent covariates			
White (vs African-American)	1,512	0.79	
Age at risk entry	1,512	26.37	4.61
Time-dependent covariates			
Is a father in:			
1968-72	15,038	0.03	0.18
1973-77	15,038	0.10	0.30
1978-82	15,038	0.18	0.38
1983-87	15,038	0.24	0.42
1988-92	15,038	0.25	0.44
1993-97	15,038	0.20	0.40
Father has some years of college	15,038	0.56	0.50
Father's labor income ($1000's)	15,038	24.70	23.13
Father's weekly employment hours	15,038	44.27	12.11
Mother has some years of college	15,038	0.49	0.50
Mother's labor income ($1000's)	15,038	7.20	9.86
Mother's weekly employment hours	15,038	22.44	18.51

Family receives AFDC	15,038	0.017	0.129

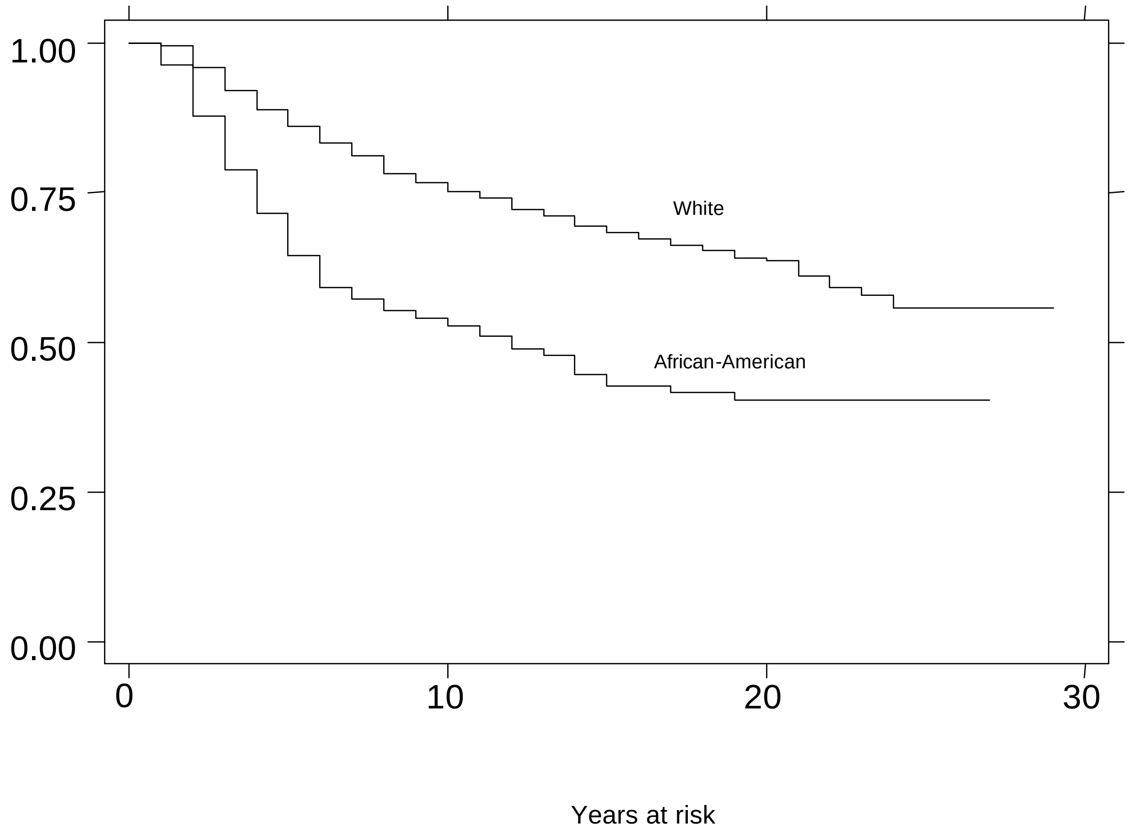
As a descriptive preliminary to our analyses, Table 2 shows the proportions of fathers who experience nonresidence in each of the six periods. These numbers are simply proportions observed to experience nonresidence in a given period; they do not account for censoring. The same fathers can be present in multiple periods. We see that the proportion increases only mildly during the 1970s, then rises substantially in the 1980s, and stabilizes during the 1990s. Our multivariate results are consistent with this basic pattern.

TABLE 2. PROPORTIONS OF FATHERS EXPERIENCING NONRESIDENCE BY PERIOD

Period	*N*	Proportion
1968-72	190	0.032
1973-77	438	0.039
1978-82	676	0.068
1983-87	908	0.067
1988-92	966	0.065
1993-96	887	0.060

Figure 1 anticipates our multivariate findings regarding the race differences in the hazard of nonresidence. In this figure, the horizontal axis is the number of years of risk, regardless of which calendar year marks the entry into risk for a particular father. The vertical axis is the proportion of fathers who remain coresident after a given number of years at risk. We see that African-American fathers are less likely than white fathers to remain coresident, or more likely to experience nonresidence, after a given number of years at risk. For example, about 50 percent of African-American fathers experience nonresidence after 10 years of coresidence, while about 25 percent of white fathers do so.

Figure 1: Proportion of whites and African-Americans remaining coresident, by number of years at risk



Period effects on the hazard

Our multivariate analyses are shown in Table 3. We estimate the hazard of nonresidence faced by a father in a given period by using indicator variables for five-year groups of years marking the beginning of spells. The reference group is the first period, which includes the spells that begin in 1968 through 1972. Our models show that the hazard of nonresidence in 1993-96, the final period, is more than twice the initial hazard in the period 1968-72. The basic pattern of period effects, with the increase in hazard concentrated in the 1980s, is stable across the multivariate models in Table 3.

TABLE 3. RELATIVE PERIOD HAZARDS OF NONRESIDENCE, BY INITIAL PERIOD

		Final period									
		1968-72	1973-77	1978-82		1983-87		1988-92		1993-96	
	1968-72	1.00	0.71	1.18		1.64		2.00	*	2.32	**
	1973-77		1.00	1.65	*	2.30	***	2.80	***	3.25	***
Initial period	1978-82			1.00		1.39	*	1.70	***	1.97	***
	1983-87					1.00		1.22		1.41	*
	1988-92							1.00		1.16	
	1993-96									1.00	

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

To highlight the period effects, we present the hazard ratios for successive periods relative to the hazard in a particular period. The results, shown in Table 4, are broadly consistent with our descriptive calculations reported in Table 2. The cells in each row of Table 4 display the hazard ratios for successive periods relative to the hazard, fixed at the value 1, in a given initial period. The significance levels of these ratios are shown alongside. (The hazard ratios are calculated from Model 2 in Table 3.) For example, the first row shows the hazards for each period following the first period, 1968-72, relative to the hazard of 1 for that period. We see that men who are fathers in 1993-96 experience nonresidence at more than twice the rate of men who are fathers in 1968-72.[6]

6 In separate models, not shown, we investigate the effects of year of risk entry on the hazard of nonresidence in addition to the period effects reported here. To do this we add an indicator variable that distinguishes men who become fathers for the first time in the first half of the thirty-year span, 1968-82, from the men who enter risk in the second half, 1983-96. There is no significant difference between the hazards of these two groups of men.

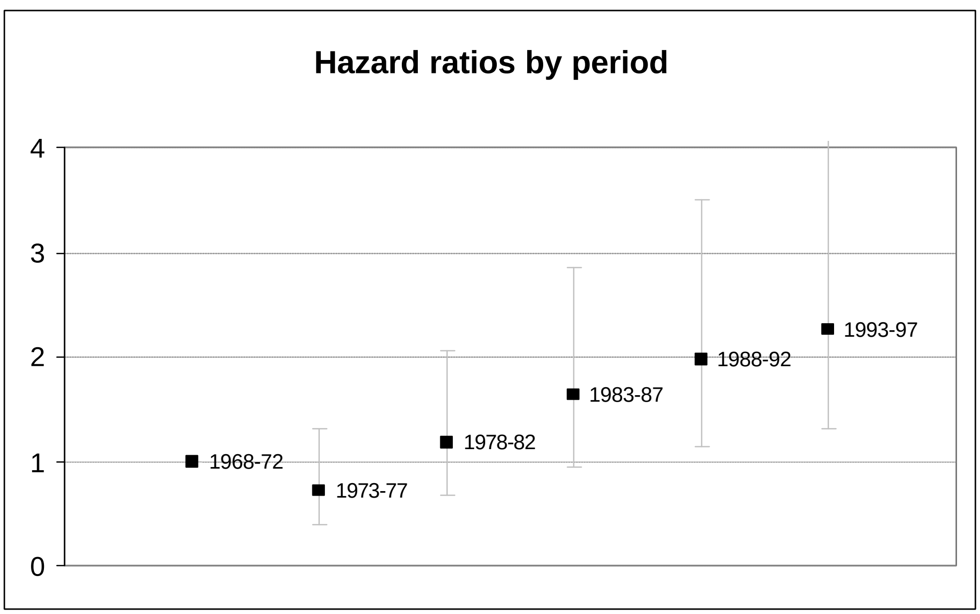
TABLE 4. COX REGRESSION MODELS OF THE HAZARD OF NONRESIDENCE

	Model 1			Model 2		
	b	exp(b)		b	exp(b)	
Is a father in:						
1968-72	Ref.			Ref.		
1973-77	-0.205	0.815		-0.337	0.714	
1978-82	0.269	1.308		0.164	1.178	
1983-87	0.646	1.908	*	0.496	1.641	
1988-92	0.652	1.919	*	0.694	2.003	*
1993-97	0.782	2.186	**	0.843	2.324	**
Age at risk entry	-0.088	0.916	***	-0.060	0.942	***
Age of youngest child				0.175	1.191	***
Number of children				-0.004	0.996	
Father's characteristics						
White				-0.682	0.506	***
Has some college				0.110	1.116	
Labor income ($1000's)				-0.022	0.979	***
Employment hours				-0.001	0.999	
Mother's characteristics						
Has some college				0.049	1.051	
labor income ($1000's)				-0.098	0.907	***
Employment hours				-0.068	0.934	***
Family receives AFDC				-1.672	0.188	***
Log likelihood	-3254.64			-2847.71		
LR Chi-Square (*df*)	80.64 (6)			893.94 (16)		
No. person-years	15038			15031		
No. persons	1512			1512		

· *p < 0.05; ** p < 0.01; *** p < 0.001*

The hazard has not increased in a smooth or linear fashion between the first and last periods, however. Rather, we discern three distinct phases in the trend: an initial period of stability, followed by a sharp increase during the 1980s, and then another period of relative stability in the last stage. In the first row of Table 4, we see that the hazard in 1973-77 is actually lower than it is in 1968-72, though this decrease is not statistically significant. The second row shows that the hazard in 1978-82 is 65 percent higher than it is in 1973-77, and this increase is significant at the 5 percent level. It is evident in the third row that the hazard in 1983-87 is almost 40 percent higher than it is in 1978-82; that increase is also significant. The subsequent increases in the hazard are not significant. While the hazard in 1988-92 is 22 percent greater than it is in 1983-87, and the hazard in the final period is 16 percent greater than it is in 1988-92, neither of these increases is statistically significant. Thus most of the increase in the hazard occurs during the middle decade, with relative stability in the two decades flanking it. Table 4 also allows us to compare hazards in non-consecutive periods. For example, the hazard in 1988-92 is 70 percent higher than it is a decade earlier in 1978-82. Likewise the hazard in the final period, 1993-96, is 40 percent higher than it is in the previous decade, 1983-87. The trend in the hazard of nonresidence is depicted visually in Fig 2, with the hazard in 1968-72 set to 1.

Figure 2: Trend in hazard ratios, 1968-1997



The covariates in our models serve as control variables. Both models in Table 3 show that men who enter fatherhood at older ages are less likely to experience nonresidence. Model 2 shows that for small increases in age, the hazard of nonresidence decreases by a little less than 6 percent for each year. The hazard for men who are 30 when they have their first child is almost 55 percent lower than it is for men who are 20 when they enter the risk set. The results in Model 2 also show that the hazard of nonresidence for fathers increases with the age of their children. Every additional year of age of the youngest child increases the hazard by about 19 percent, for small age increments; the hazard of fathers whose youngest child is 10 years old is about four times that of fathers whose youngest child is 2.

Model 2 reveals the effects of the main socioeconomic covariates. Chief among these is race: the hazard of nonresidence for white fathers is almost 50 percent smaller than it is for African-American fathers. Having some years of college education, by contrast, has no impact on the hazard of nonresidence. Earnings do, however. The hazard decreases by almost 2 percent for each additional $1,000 of earnings; the hazard for fathers who earn $30,000 is 20 percent lower than it is for fathers who make $20,000. Fathers' employment hours do not affect their hazard.

Certain characteristics of mothers also affect the fathers' hazard of nonresidence. While their education does not have a significant effect, their earnings and employment hours do. The more they earn and the more time they spend on paid work, the lower the fathers' hazard of nonresidence. An additional $1,000 of mothers' earnings reduces the hazard by a little over 9 percent, a larger effect than the 2 percent resulting from the same increment in fathers' earnings. This means that the hazard for fathers whose wives or partners earn $30,000 is 37 percent lower than it is for fathers whose wives or partners earn $20,000. Further, unlike fathers' employment hours, those of mothers have a large dampening effect on the hazard, with every additional hour reducing the hazard by a little over 6 percent. The hazard for fathers whose partners work 40 hours a week is 50 percent lower than it is for fathers whose partners do not work outside the home. Finally, the hazard for fathers in families that receive AFDC support is more than 80 percent lower than it is for fathers in families that do not.[7]

Our results for earnings and employment hours lend support to the ‘income effect’ hypothesis of marital dissolution, namely that wives’ earnings reduce marital instability. The principal contending argument is the ‘independence’ hypothesis: women’s earnings and employment increase the likelihood of marital disruption. The research on these competing hypotheses has yielded mixed results (Brines and Joyner 1999; Dechter 1992; Greenstein 1990; Ono 1998; Oppenheimer 1997; Ruggles 1997). We tested the independence hypothesis by adding the ratio of women’s earnings to father’s earnings, both with and without the absolute earnings levels. Our results were the same: mothers’ earnings, whether relative or absolute, reduce the hazard of nonresidence for fathers. Since fathers’ earnings and AFDC support also have negative effects on the hazard, we conclude that greater family income, whatever its source, reduces the hazard of nonresidence for fathers. We obtain the same results with alternative specifications for income and employment, using categorical rather than continuous variables for both; accordingly we retain the continuous variables for parsimony.

DISCUSSION

Two distinct characterizations of fatherhood have evolved in the scholarly literature and in popular culture, the ‘deadbeat dad’ on the one hand, and the nurturing father on the other. Both begin with a concern for children’s well-being, and both recognize that trends in fatherhood are a critical part of changes in family behavior over the last few decades. At the same time, the exact nature of these trends has not been thoroughly documented because of the retrospective reports of fertility in most surveys. These result in undercounts and nonreporting, making it difficult to identify and describe nonresident fathers.

Our primary aim in this study is to use panel-updated fertility reports and supply the public and

7 There is no collinearity among the variables for father's education, earnings and employment hours. Separate models that introduce each of these in turn show that the coefficient of one variable is not affected substantially by the presence of the other variables. There is some collinearity between the mother's earnings and employment hours. In a model that excludes employment hours, each additional $1,000 of mother's earnings reduces the father's hazard by 30 percent. After we add her employment hours, that earnings effect is reduced to the 9 percent discussed earlier. Still, it remains large and significant, and we choose to include both variables because each

scholarly discussion of nonresidential fatherhood with some basic empirical information on trends in fathers' nonresidence. Our period models show that biological, coresidential fathers experience nonresidence at higher rates in later years during the period 1968-97 than they do earlier in that period. At the same time, we find that this rise in nonresidential fatherhood among men who once lived with their children has not been steady or linear. Instead, it seems to have occurred in three stages: (i) a period of flat growth for most of the 1970s; (ii) a decade of sustained increase starting in the late 1970s and continuing through the late 1980s; and (iii) a leveling out during the 1990s. Thus, references to the trend in nonresident fatherhood ought to acknowledge that the increase has not been constant.

There are some problems in our data that could detract from the generalizability of our results. We could not avoid altogether the use of retrospective fertility data, though the disadvantages are offset by our use of panel-updated data for a substantial part of the study period. Additionally, the PSID loses an increasing number of fathers due to attrition over the years. To the extent that these losses are related to, or caused by, the event of nonresidence, our estimates of the hazard of nonresidence in later periods could be biased downward. However, our investigation of this possibility shows that it is not a major concern.[8]

Our results are broadly consistent with Goldstein's (1999) finding that the hazards of divorce have leveled off in recent years. They also support Casper and Bianchi's (forthcoming) finding that the 1990s have been a time of slower change and of leveling in family trends generally. More generally, our findings suggest that even brief historical periods may contain considerable fluctuation, and thus may be better understood and more accurately described by change rather than continuity.

In addition to analyzing the period effects on the hazard of nonresidence, we examine the effects of race and income. We find that the hazard for African-American fathers is about 50 percent greater than

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of them has a substantial and significant effect on the father's hazard of nonresidence.

8 We find that the period trend in the hazard is not affected by survey attrition. To examine the impact of attrition, we assume that the attrited fathers experience nonresidence, in each period, at the same rate as do the fathers who remain in the survey. Then we re-estimate our models, and find that our period coefficients change only slightly. Further, we explore the possibility that the fathers who are lost to the survey are more likely to experience nonresidence than the fathers who remain. We assume that in each period, the

it is for white fathers, and that every additional $10,000 of income reduces the hazard by 20 percent. Our study thus adds to previous research that links economic hardship to family instability. Those concerned with the formation and maintenance of two-parent families should recognize that it is easier for fathers to remain with their children if their incomes are higher. The additional finding that AFDC assistance reduces the hazard of nonresidence corroborates this result. It is also important to note that fathers are less likely to leave their children when the mothers earn higher incomes and work a greater number of hours. This finding adds to the growing body of evidence that women's income may have a stabilizing rather than disruptive effect on family patterns.

Our findings are grounds for both concern and for cautious optimism. On the one hand, the hazard of nonresidence for fathers has more than doubled over the last three decades. On the other, this increase has not occurred in the steady, implacable fashion that would suggest inevitability. Rather, it seems to be characterized by periods of stability as well as change. Future research should focus on explaining these rhythms in the individual-level period effects, so that we can understand why the hazard has increased in some periods and not in others.

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attrited fathers experience nonresidence at substantially higher rates than do the fathers who remain in the survey. Our period coefficients are not affected greatly under these conditions, and the basic trend is unchanged.

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