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Nonresident Fatherhood and Adolescent Sexual Behavior: A Comparison of Siblings Approach

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Although voluminous research has linked nonresident fatherhood to riskier sexual behavior in adolescence, including earlier sexual debut, neither the causality of that link nor the mechanism accounting for it has been well-established. Using data from the National Longitudinal Survey of Youth, 1979—the Young Adult Survey (CNLSY-YA), the present study addresses both questions by comparing the sexual development of siblings discordant for age at father departure from the home and examining results across behavioral (age at first intercourse), biological (pubertal timing), and cognitive (attitudes about childbearing and marriage) sexual outcomes ($N = 5{,}542$). Findings indicate that nonresident fatherhood, beginning either at birth or during middle childhood, leads to an earlier sexual debut for girls, but not for boys, an effect likely explained by weak parental monitoring rather than an accelerated reproductive strategy.

Keywords: nonresident fatherhood, sexual behavior, comparison of siblings, puberty, adolescence

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There exists a well-established association between nonresident fatherhood—as the result of nonmarital childbirth or divorce—and risky sexual behavior during adolescence. Teenagers who have experienced nonresident fatherhood initiate sexual behavior at an earlier age (D'Onofrio et al., 2006; Ellis et al., 2003; James, Ellis, Schlomer, & Garber, 2012), engage in riskier sexual behavior (James et al., 2012; Kiernan & Hobcraft, 1997), and have more sexual partners (Quinlan, 2003) than those who live with both biological parents until adolescence. These outcomes, in turn, elevate the risk of sexually transmitted disease and teenage pregnancy (Hofferth & Hayes, 1987; O'Donnell, O'Donnell, & Stueve, 2001), and the likelihood of nonmarital childbirth and family instability (Gee & Rhodes, 2003; Teachman, 2002). The apparent salience of nonresident fatherhood for sexual behavior, fertility, and family formation in the next generation validates public concern over the precipitous rise in nonresident fatherhood over the past half century (U.S. Census Bureau, 2006; Ventura, 2009) and suggests public policy efforts should be directed at reducing its incidence or ameliorating its effects.

Before advocating or enacting such policies, however, research is needed to answer two outstanding questions. First, is the association between nonresident fatherhood and risky sexual behavior causal? It is neither possible nor ethical to randomly assign children to family structure experiences. Thus, family level factors, both genetic and environmental, that select parents into family disruption may influence adolescent sexual behavior and in doing so drive the documented nonresident fatherhood–sexual behavior links. If we can establish that nonresident fatherhood–sexual behavior associations are at least plausibly causal, research must still address a second key question: by what mechanism does nonresident fatherhood lead to riskier sex? Only by understanding *how* nonresident fatherhood shapes sexual development can we understand fathers' role in children's sexual development and identify policy approaches to ameliorating the impact of their nonresidence. The literature on these associations offers three broad theories. The *paternal investment theory* (PIT), an extension of Belsky, Steinberg, and Draper's (1991) influential *psychosocial acceleration theory,* posits that early father departure signals to offspring that paternal investment is not essential to reproduction and modifies their neurophysiologic and motivational systems to speed pubertal maturation, accelerate sexual debut, and orient them toward weak (uncommitted or unstable) pair bonds (Ellis, 2004). By contrast, socialization theory posits that father absence models sexual attitudes favoring weak commitments and, thus, earlier, riskier sexual behavior (Amato & DeBoer, 2001). Finally, parental monitoring or social control theory holds that father absence facilitates earlier, riskier sexual behavior via reduced parental supervision rather than attitudes (or biology) per se (Hogan & Kitagawa, 1985; Newcomer & Udry, 1987).

The present study addresses both questions by comparing the sexual development of siblings. Comparing siblings within families, rather than unrelated youth across families, reduces the influence of unobserved genetic and environmental risks that vary between families, better isolating any causal effect of nonresident fatherhood on sexual behavior. Although siblings typically share the experience of nonresident fatherhood, they also differ in age

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(among nontwins), thus I can compare the effect of nonresident fatherhood among siblings by distinguishing effects by child age at the time of disruption. Comparing the effects across child age also addresses the question of mechanism: two of the three theories described above suggest distinct patterns of effects by age at father departure. I also investigate mechanisms by comparing effects of nonresident fatherhood across different sexual domains. All three theories suggest nonresident fatherhood should impact sexual behavior; however, only the PIT suggests this process includes pubertal timing, and only the PIT and socialization theory suggest it operates through sexual attitudes. Because each theory implies a unique pathway, any variation in the impact of nonresident fatherhood by child age or sexual domain could pinpoint the active mechanism.

Nonresident Fatherhood and Risky Sexual Behavior: A Causal Link?

Although the link between nonresident fatherhood and risky sexual behavior has been well-established, the causality of that link has not. Genes passed from parents to children, such as those for early puberty (Newcomer & Udry, 1984; Rowe, 2002), impulsive, externalizing behavior (Raine, 2008; Verweij, Zietsch, Bailey, & Martin, 2009), and sexual behavior itself (Rodgers, Rowe, & Buster, 1999; Rowe, 2002) may trigger early and risky sexual behavior, which in turn predicts unstable relationships and nonresident fatherhood. Environmental factors such as low income, neighborhood disadvantage, and other familial stressors also covary with family disruption and risky sexual outcomes (Browning, Leventhal, & Brooks-Gunn, 2005; Kirby, 2002; Moore & Chase-Lansdale, 2001) and could confound associations. In short, familial factors that select families into disruption, both genetic and environmental, may also trigger earlier and risky sexual behavior, thus inducing a spurious link between the two.

The typical approach to minimizing influence of selection is to control for a robust set of environmental factors that covary with nonresident fatherhood and predict sexual behavior. This approach has two problems: (a) it cannot eliminate the influence of all possible confounds because some factors are unmeasured and others likely unmeasurable; (b) it does not address genetic confounds. A more rigorous way to estimate whether links between family experiences and child outcomes are plausibly causal is to use a quasi-experimental design in which relatives discordant for the family experience of interest are compared rather than unrelated youth. For example, it is common in public policy and economic research, and more recently practiced in psychology (Lahey & D'Onofrio, 2010), to compare cousins who differ on a particular family experience (Coyne, Långström, Rickert, Lichtenstein, & D'Onofrio, 2013; Geronimus, Korenman, & Hillemeier, 1994). This design, sometimes called the Children of Siblings (COS) approach, reduces the influence of genetic risks that could bias associations because cousins share either 12.5% or 25% of their genetic makeup depending on whether they are the offspring of full siblings (or fraternal twins) or monozygotic twins. Presumably, cousins also experience more similar home environments than unrelated youth because their mothers were raised together, and this similarity minimizes the influence of unobserved environmental confounds.

The few studies that have used the COS approach to examine links between nonresident fatherhood and risky sexual behavior have yielded conflicting findings. D'Onofrio and colleagues (2006) found that father absence predicted earlier sexual debut, with effect sizes comparable with those in between-family studies. However, Mendle and colleagues (2009) reached the opposite conclusion, with regard to age at first intercourse, when comparing children of sisters in the National Longitudinal Survey of Youth. Their findings may conflict because D'Onofrio used a high-risk Australian sample and Mendle a U.S. sample. Both studies also used relatively small samples of discordant pairs, substantially limiting their statistical power, as well as variability on nonresident fatherhood and sexual behavior, limitations that may have undermined their ability to reliably estimate associations across cousins (Coyne et al., 2013).

More importantly, the COS approach leaves open the possibility that genetic and environmental differences between sisters, and between cousins' unrelated fathers, may confound the association between nonresident fatherhood and risky sexual behavior. For instance, sisters who chose different reproductive pathways (and partners) may differ in their genes or experiences in ways that influence their children's sexual behavior independent of family structure (East & Jacobson, 2000). Moreover, cousins' unrelated fathers might pass on a genetic predisposition for risky sexual behavior, as well as father departure, which drives a spurious link between the two. An alternate approach, the differential sibling-exposure design (Tither & Ellis, 2008), compares siblings discordant for age at father departure on sexual outcomes. When full siblings are compared, this approach eliminates the influence of genetic risk because genetic differences between siblings are effectively randomized during meiosis (Ellis, Schlomer, Tilley, & Butler, 2012; Lahey & D'Onofrio, 2010). It also eliminates the influence of environmental confounds that differ between families. When half siblings are compared, the sibling comparison approach allows for fathers' genetic differences to potentially bias associations, although this comparison still permits less genetic confounding than standard COS designs because half siblings share 25% of their genetic makeup rather than the typical 12.5% among cousins. Moreover, although the home environments of half siblings differ more than those of full siblings, their home environments are more similar than those of cousins raised by different mothers, even mothers who are monozygotic twins.

The only studies that have employed the differential sibling-exposure method to address the question of father absence found that full siblings with greater exposure to nonresident fatherhood reached menarche at an earlier age and exhibited riskier sexual behavior (measured as number of sexual partners and high-risk sexual activity; Ellis et al., 2012; Tither & Ellis, 2008). However, the effect on sexual behavior emerged only when the quality of fathers' parenting was relatively high, which jibes with other research that finds fathers' coresidency benefits child development only when fathers have the socioemotional and socioeconomic resources necessary to enhance children's home environments (Jaffee, Moffitt, Caspi, & Taylor, 2003; Ryan, 2012). It would be premature to conclude, however, that no average effect exists on the basis of this study. First, like the studies using COS designs, Ellis and colleagues (2012) had a relatively small sample and thus limited power—and perhaps limited variability on sexual behavior—to detect significant within-family effects. More importantly, the study could not by design compare effects by developmental period or estimate the specific effect of early father departure, for father departure occurred during middle or late childhood for sisters in their study. For reasons explained below, it is plausible that the effect of father departure varies by age and that effects are largest when children are very young. Thus, although the studies using COS and differential sibling exposure designs made significant methodological advances, it is still unclear whether within-family comparisons reveal an average impact of nonresident fatherhood on sexual development.

Nonresident Fatherhood and Risky Sexual Behavior: Pathways of Influence

If links between nonresident fatherhood and adolescent sexual outcomes are determined to be plausibly causal, understanding *how* nonresident fatherhood shapes sexual development could help identify appropriate policy interventions. One way to investigate mechanisms is to compare impacts across child age at father departure. According to PIT, father departure before age 5, when children are forming fundamental attachment relationships with caregivers, signals that paternal investment of emotional and other resources are unavailable or unpredictable. Human beings have evolved to respond in these environments, the theory maintains, to reach puberty earlier, initiate sexual behavior earlier, and develop weaker pair bonds to maximize family size and thus the chance some offspring will reproduce in the absence of strong parental investment (Belsky et al., 1991; Boyce & Ellis, 2005). Thus, according to PIT, father departure before age 5 should impact sexual development more than departure later on (Ellis, 2004). Socialization theory too implies that father departure would have a stronger impact on sexual behavior when it occurs early on because younger children spend more time exposed to the socializing influence. In contrast, monitoring theory suggests that father departure any time before puberty should equally impact sexual development because children would reach adolescence without a monitoring father. Although comparing effects by age cannot distinguish cleanly between PIT and socialization, it could highlight the potential role of monitoring.

Existing research comparing the sexual outcomes of unrelated adolescents offers conflicting findings about timing. Some find no difference between early versus later disruption on sexual outcomes (Teachman, 2002, 2003), others find that later transitions have a stronger effect on sexual behavior (Cavanagh, Crissey, & Raley, 2008), whereas others find only earlier father departure leads to earlier sexual behavior (Quinlan, 2003). Previous studies using within-family comparisons have not rigorously investigated timing effects, perhaps because they have not had large enough samples to differentiate by child age.

Another way to investigate mechanisms is to compare impacts of nonresident fatherhood by sexual domain. PIT maintains that father absence should impact sexual development across biological (earlier puberty), behavioral (sex and reproduction outside committed pair bonds), and cognitive (attitudes favoring sex and reproduction outside committed, stable pair bonds) domains (Ellis, 2004). A pattern of results in which nonresident fatherhood impacts outcomes across all three domains would support the mechanism PIT hypothesizes—an accelerated and less restrictive reproductive strategy. Socialization theory, however, suggests nonresident fatherhood impacts sexual attitudes and behavior but does not posit early pubertal timing as a related outcome. Parental monitoring acknowledges nonresident fatherhood impacts sexual behavior but does not imply impacts on sexual attitudes or pubertal timing. Thus, the pattern of findings across biological, behavioral, and cognitive outcomes could illuminate which theory best explains the link between nonresident fatherhood and risky sexual behavior.

Although individual studies have examined these outcomes, none have used this comparison-across-outcome approach to identify mechanisms. The larger literature has focused either on pubertal timing (Belsky et al., 2007; Tither & Ellis, 2008) or broad life course indicators such as age at first intercourse and first birth (Belsky, Schlomer, & Ellis, 2012; Ellis & Garber, 2000; James et al., 2012; Kiernan & Hobcraft, 1997; Moffitt, Caspi, Belsky, & Silva, 1992; Moore & Chase-Lansdale, 2001). Only one study has examined sexual attitudes in relation to nonresident fatherhood (Hoier, 2003) even though they often precede risky sex. This study found that single motherhood was associated with less restrictive sexual attitudes. Moreover, only two of the studies that have examined pubertal timing, and none of the studies examining attitudes, have used a within-family comparison approach, even though both outcomes could be influenced by the same genetic and environmental confounds that plague the association between nonresident fatherhood and sexual behavior. Finally, those that have used a within-family approach to examine pubertal timing offer conflicting findings (Mendle et al., 2006; Tither & Ellis, 2008).

Nonresident Fatherhood and Risky Sexual Behavior: Gender Differences

Both theory and research on the link between nonresident fatherhood and adolescent sexual behavior suggests associations may differ by child gender. Boys tend to have fewer intimate friendships than girls during adolescence (Maccoby, 1998), so they may turn to romantic or sexual relationships for support more readily in response to a family stressor like nonresident fatherhood. Boys may also engage in risky sex more readily in response to father absence than girls if fathers are stronger sexual role models for sons. Most theories, however, suggest girls' behavior should be more affected by father absence. Girls are more attuned to relationships and relationship quality than boys; in turn, their relationship skills might develop more strongly in response to family dynamics (Amato, 1993; Crockett & Randall, 2006). According to PIT specifically, girls have evolved to be more attuned to paternal investments of emotional resources during early childhood because females depend more on familial resources during pregnancy and child rearing than males (Gangestad & Simpson, 2000; Jackson & Ellis, 2009). This theory, therefore, suggests nonresident fatherhood should have unique effects on girls' sexual behavior because it has unique implications for their reproductive strategy.

Evidence on gender differences in the association between nonresident fatherhood and sexual behavior is mixed. Some research finds that family structure instability is more strongly associated with boys' sexual behavior (Cavanagh et al., 2008) or impacts boys and girls similarly (Ryan, Franzetta, Schelar, & Manlove, 2009). However, most studies find that girls respond more to family disruption in terms of sexual behavior than boys (Davis &

Friel, 2001; James et al., 2012; Thornton, 1991). Recently, James and colleagues (2012) found that nonresident fatherhood had a direct effect on girls' but not boys' sexual risk taking, an effect mediated through accelerated pubertal timing (age at menarche). This study jibes with other work looking specifically at pubertal timing that finds girls, but not boys, reach puberty earlier when fathers are nonresident, a difference not entirely attributable to less precise measurement of boys' pubertal timing (Belsky et al., 2007; James et al., 2012). In sum, there is some evidence that girls' respond more to father absence than boys in terms of sexual behavior and that a stronger biological sensitivity to early paternal investment may account for the difference.

Present Study

The present study addresses two outstanding questions about the association between nonresident fatherhood and adolescent sexual behavior: (a) are the documented links plausibly causal; and (b) if so, what mechanism likely accounts for those links. It addresses the first question by comparing siblings discordant for the experience of nonresident fatherhood. Data are drawn from a diverse, national dataset that offers a larger sample of related pairs than any previous study on this topic using a within-family approach, thus maximizing power to detect within-family effects of father absence. To that end, main analyses include both full and half siblings (all born to the same mother) to retain a sufficient number of sibling pairs. To estimate the bias introduced by including half siblings, sensitivity analyses will be run on the smaller, full sibling sample. The question of mechanism is addressed by comparing impacts of father absence by timing and across behavioral, biological, and cognitive outcomes. Special attention is paid to differences in associations by child gender. By addressing both *how* nonresident fatherhood may impact sexual development, as well as *whether* it does, the study aims to illuminate the unique role fathers may play in children's sexual development and approaches to alleviating the potential effects of their nonresidence.

Method

Data and Sample

Data are drawn from the Children of the National Longitudinal Survey of Youth, 1979—the Young Adults Survey (CNLSY-YA). The National Longitudinal Survey of Youth, 1979 (NLSY79) is a nationally representative sample of 12,686 young men and women who were 14–22 years old when they were first surveyed in 1979 and are currently interviewed on a biennial basis. Since 1986, the biological children of the NLSY79 mothers have been independently followed. Starting in 1994, children who reached the age of 15 by the end of the survey year, called "youth adults" (YA), were given complete personal interviews akin to those given to their mothers during late adolescence and into adulthood. The most recent data used in the present study were collected in 2008 when YA respondents were between 15 and 37 years old (Center for Human Resource Research, 2009). Because most female respondents had more than one child, the NLSY79 contains a large number of sibling pairs, making it ideal for a within-family approach. The analytic sample was limited to all youth with at least one YA interview, at least one interviewed sibling, and data on age at father departure from the home (N = 6,141 individual YAs drawn from 2,330 nuclear families). The sample was further limited to youth with valid data on sexual outcomes, which ranged across measures (see below).

Descriptive statistics for all dependent variables and all covariates by age at father departure from the home are presented in Table 1, using the sample with valid data on the key dependent variable, age at first intercourse (N = 5,542; descriptive statistics did not vary substantively in analytic samples for the other dependent variables). The analytic sample is disadvantaged relative to national norms because youth who were at least 15 years old in 2008, the criterion for inclusion in the YA study as of 2008, were by design disproportionately born to young mothers. As a result, over a third of children in the analytic sample were African

Table 1
Descriptive Statistics by Age at Father Departure

	Full sample	Father always absent	Father left early (0–5)	Father left late (6–13)	Father always present
Age of first intercourse	15.65 (2.15)	15.00 (1.98)$_a$	15.44 (2.20)$_b$	15.43 (2.12)$_b$	16.16 (2.10)$_c$
Had sex before censoring (%)	74.09	86.53	80.34	81.09	64.77
Age at menarche (girls only)	12.24 (1.37)	11.95 (1.28)$_a$	12.21 (1.59)$_{ab}$	12.27 (1.37)$_b$	12.37 (1.30)$_b$
Ideal age at childbirth	25.16 (4.39)	24.14 (5.11)$_a$	25.08 (4.50)$_b$	25.35 (4.50)$_b$	25.48 (3.96)$_b$
Ideal marriage < Ideal childbirth	79.31	60.27	74.82	80.43	86.81
Male (%)	50.90	52.08	48.78	49.10	51.49
Family has half siblings (%)	36.00	83.03	56.54	31.68	13.60
Mother's education level (%)					
Less than high school	37.60	60.23	42.30	40.59	26.74
High school degree	36.11	28.42	38.89	37.03	37.50
Some college	18.15	11.03	15.62	17.82	21.71
College graduate	8.14	0.32	3.19	4.55	14.05
Race of child (%)					
Non-Hispanic/Non-African American	42.08	11.45	35.92	42.87	54.87
African American	34.55	74.34	40.70	30.10	19.90
Hispanic	23.37	14.21	23.38	27.03	25.23
Mother's age at first birth	20.99 (3.81)	18.88 (2.93)	20.11 (3.33)	20.54 (3.41)	22.23 (3.94)
Average household income (ln)	9.86 (0.69)	9.31 (0.58)	9.66 (0.61)	9.78 (0.60)	10.15 (0.63)

Note. N = 5,542 for age at first intercourse; N = 1,822 for age at menarche; N = 4,302 and 4,017 for ideal childbirth age and ideal marriage age before childbirth age, respectively. Means for sexual outcomes with different subscripted letters are significantly different at $p < .05$.

American, over a third had half siblings, and over a third of mothers had less than a high school degree at the time of their first birth.

Measures

Age at father departure. Information on timing of father departure from the home was gathered from various sources in the NLSY79 and CNLSY-YA. Youth were asked during their YA interview if they lived with their biological father and, if not, when they last lived with him. If youth reported on age at father departure in multiple interviews, responses were drawn from the earliest wave. Mothers also reported whether the child lived with his or her biological father in each CNLSY-YA mother interview prior to the youth entering the YA study. Finally, in the mothers' main CNLSY-YA interview, she was asked about her marriage and cohabitation history. This information was used to create complete marriage and cohabitation histories for mothers.

Information from these three sources on age at father departure was combined in the following way. First, for each report, youth were divided into four exclusive groups: father always present (coresiding with both parents through age 13); father never present (father left before child was born or father and mother never lived together); father left between birth and age 5; and father left between ages 6 and 13. Although these groupings are less precise than a quantitative measure, they allow for specific nonmonotonic patterns in the association between age at father departure and sexual outcomes to emerge that would support specific mechanisms in question. For example, the PIT posits that father departure in the first 5 years of life exerts a stronger influence on sexual outcomes than later departure. A quantitative measure of age at father departure that estimates the effect of one year on sexual outcomes would not capture this distinction as clearly as the categorical groupings, even if nonlinear forms of the quantitative measure were included such as a quadratic term. Those whose fathers departed during the teenage years were categorized as father always present because for those youth puberty and sexual intercourse would be increasingly likely to have occurred before father departure. Next, the youth and mother report from the CNLSY-YA, and the interviews were compared. If the youth and mother reported the same age period for father departure, which they did in 75% of cases, that age period was used. For all remaining cases, information based on the mothers' marriage and cohabitation history was used. This resulting four-level variable was recoded into three indicator variables for father always absent, father left between birth and age 5, and father left between ages 6 and 13, with father always present as the reference.

Within-family deviation in age at father departure. In order to estimate the within-family effect of father departure at different child ages, a child-level deviation from the family's average for each father absence indicator was computed. First, within-family averages on each father absence indicator were generated. For example, if there were two children in a family and the parents separated when one child was 2 and the other was 6 years old, the family average for father left between birth and age 5 would be .5, and the family average for father left between ages 6 and 13 would be .5. Second, child-level deviations from the family average were calculated by subtracting the average from each child's score on each indicator. So, the 2-year-old would have a deviation score of .5 for father left between birth and age 5, whereas the 6-year-old would have a deviation score of −.5. For father left between ages 6 and 13, the deviation scores would be reversed. In this way, the child-level deviations for each father absence category operate as within-family dummy variables. Siblings concordant for age at father departure, including those who never experienced father departure, receive 0 for all child-level deviations. In this formulation, the within-family deviations represent manual constructions of family level "fixed effects" commonly used in econometrics (e.g., Geronimus et al., 1994); it is a standard coding scheme used for estimating within-family effects in developmental research (see Mendle et al., 2009).

Although most siblings in the analytic sample had the same nonresident fatherhood experience, there was adequate discordance among siblings to estimate within-family effects. The child-level deviations are calculated to reflect whether the child differs from any siblings on each father absence category. However, because the reference group in models is father always present, siblings who experienced father departure at each age are compared with siblings who had fathers always present. Thus, it is most important to consider the number of siblings who differ in this way. Among the 5,542 in the age at first intercourse analyses, 250 of youth who had their fathers always present had a sibling whose father left between ages 6 and 13, 167 had a sibling whose father who left between birth and age 5, and 135 had a sibling whose father was always absent. Not surprisingly, families in the latter group had the largest proportion of half siblings (86% had a half sibling in the family vs. 32% for those with a sibling whose father left after age 6).

Sexual outcomes.

Sexual behavior. The CNLSY-YA asked if youth had ever had sexual intercourse and, if so, age in years at first intercourse. Six percent of youth reported on age at first intercourse in more than one interview; in these cases, the earliest report was used to minimize telescoping bias. The age at first intercourse is used as the indicator of risky sexual behavior, with earlier age indicating riskier behavior. The average age at first intercourse was 15.7 (SD = 2.15); however, 26% of the analytic sample had not had sexual intercourse by the time of their last YA interview.

Pubertal timing. The CNLSY-YA asked girls if they ever had a menstrual period and, if so, at what age they reached menarche. Information on age at menarche was drawn from the first interview in which the youth reported experiencing menarche (M = 12.24, SD = 1.37). The CNLSY-YA did not ask boys about their pubertal development, so analyses on pubertal timing were conducted only with girls who also had female sibling in the YA sample (N = 1,822), so that within-family deviations could be computed.

Sexual and relationship attitudes. In the CNLSY-YA, youth are asked at what ages they would ideally get married and have a first child. Analyses examined ideal age at first childbirth because an earlier ideal age of childbirth reflects an orientation toward reproduction within weaker, less stable relationships (M = 25.16, SD = 4.39). A dichotomous variable was also constructed reflecting whether the age at ideal marriage precedes age at ideal childbirth. Using this variable assumes an endorsement of marital birth, after which relationships are less likely to dissolve than after a nonmarital birth; thus, endorsing marital birth reflects orientations toward stronger, more stable relationships (79.3% reported an ideal marriage age younger than ideal childbirth age). Ideally, the

CNLSY-YA would have asked more directly about attitudes toward sexual behavior, such as endorsement of sex outside of committed relationships, infidelity, and ideal partner characteristics (see Hoier, 2003). However, because earlier age at childbirth and nonmarital birth both predict relationship dissolution and instability, as well as nonresident fatherhood, they offer adequate proxies for orientation toward paternal investment. Responses to these questions were drawn from youths' earliest interview (usually age 15).

Half sibling status. Youth were first asked about their relatedness to siblings in the YA study in 2006. They were asked whether they shared a father, did not share a father, or did not know if they shared a father with each interviewed sibling. For youth missing data on sibling relatedness, older siblings' valid responses were used (or younger siblings' if the youth had no older sibling or older siblings were missing data). If two or more siblings had valid responses, but those responses conflicted (one sibling reported they were half siblings, whereas the other reported they were full siblings), the older siblings' data was used for the family. This information was then used to determine if the youth had any half siblings in his or her family, with those who reported not knowing their relatedness to a sibling coded as having a half sibling. Six percent of youth in the analytic sample had missing data on youth reported half sibling status. In these cases, relatedness was deduced from mothers' report of biological father coresidence in the following way: if a mother reported that one sibling coresided with his or her biological father but another sibling did not in a single year, the siblings were coded as half siblings; if a mother reported that two siblings both coresided with their biological father in a single year, they were coded as full siblings. Finally, siblings were coded as twins if they had the same birth month and year (as all siblings shared a biological mother), and all twin pairs were coded as full siblings. After these deductions were made, half sibling status was still missing for 138 youth.

Covariates. Child and family level covariates exogenous to father absence were included in all models. Child-level covariates were characteristics that could vary across related children and confound within-family associations: child gender, child's birth year (to control for cohort effects and birth order), and race/ethnicity. Birth order was added as an additional covariate in supplementary models; however, it was nonsignificant when child's birth year was also included. Low birth weight status was controlled in earlier analyses but was excluded from final models due to nonsignificance. Nuclear-family covariates were parent characteristics that could vary across nuclear families, including age at mothers' first birth, mothers' education level at first birth, and the presence of half siblings (see above). Although family income is potentially endogenous to father absence, the mother's household income averaged across all interview years was entered as a measure of permanent income to control for large differences in families' socioeconomic status across nuclear families (the measure was log transformed for analyses to reduce positive skew).

Analytic Strategy

A two-level hierarchical linear model (HLM) was fit for each sexual outcome, with child-level variance at level 1 and family level variance at level 2, using STATA 12.0 (StataCorp, 2011). At level 1, each child's deviation from the family average for father always absent, father left between birth and age 5, and father left between ages 6 and 13 was entered along with child-level covariates. At level 2, family level averages on each father absence indicator were entered along with mother-level covariates. The combined level 1 and 2 model takes the following form:

$$\begin{aligned} Y_{ij} = {} & \gamma_{00} + \beta_{1j}\text{allab\_cdev}_{ij} + \gamma_{01}\text{allab\_fav}_j + \beta_{2j}\text{earab\_cdev}_{ij} \\ & + \gamma_{02}\text{earab\_fav}_j + \beta_{3j}\text{latab\_cdev}_{ij} + \gamma_{03}\text{latab\_fav}_j \\ & + \sum \beta_{qj}(\text{ChildVars})_{ij} + \sum \gamma_{0q}(\text{MotherVars})_j + u_{0j} + r_{ij} \end{aligned}$$

The variables allab_cdev_{ij}, earab_cdev_{ij}, and latab_cdev_{ij} reflect the child-level deviations from the family average; the variables with the suffix _fav_j are the analogous family level averages. With family level averages held constant, the associated coefficients for the child-level deviations (β_{1-3j}) estimate the within-family effect of father departure at different ages (Rabe-Hesketh & Skrondal, 2012). These estimates will illuminate whether associations between nonresident fatherhood and each sexual outcome are plausibly causal because they are less biased than the between family estimates by environmental and genetic differences across families. Moreover, comparing across within- and between-family estimates can illuminate the degree of bias genetic and environmental confounds introduce into nonresident fatherhood–sexual behavior associations. This modeling strategy has been used in other similar studies to estimate within-family effects of father absence (e.g., Mendle et al., 2009).

Another strength of HLM for these analyses is that it can accommodate non-normal dependent variable distributions including binary and count data (Raudenbush & Bryk, 2002). For continuous dependent variables—age at menarche (all but 8 girls had a reported menarche age in the YA sample, so right-censoring did not call for a hazard model) and ideal age at first childbirth—multilevel mixed-effects linear regression models were run with random intercepts at the family level. However, for the right-censored variable age at first intercourse, discrete multilevel hazard models were run, fit similarly to the HLMs described above. I also ran multilevel survival models on age at first intercourse with a parametric Weibull distribution, as well as a Cox regression model, and results were nearly identical to those obtained using a discrete hazard model that included a linear and quadratic time measure—this hazard model distribution has been used in other similar studies to estimate father absence associations with age at first intercourse (Mendle et al., 2009). For the dichotomous variable, ideal age at marriage precedes ideal age at childbirth, multilevel mixed-effects logit models were run, again fit similarly to the models described above.

Results

Bivariate Comparisons by Age at Father Departure

Means and standard deviations, or percentages, for all dependent variables by youth's age at father departure are reported in Table 1. Youth with fathers always absent reported a younger age at first sex than all other youth, and youth whose fathers left after birth reported younger ages at first sex than those with fathers always present. In line with the PIT, girls whose fathers were always absent or left during early childhood had an earlier age at menarche than other groups, although only differences between girls with fathers always absent and the other groups were significant. In line with socialization theory, differences in sexual attitudes by age at father departure were striking. Youth whose fathers were always absent reported a significantly lower ideal age at childbirth than all others, by over year relative to those whose fathers were always present. Moreover, only 60% of youth with fathers always absent reported an ideal age at marriage younger than their ideal age at childbirth versus 87% of youth with fathers always present.

Between Versus Within Family Estimates

Age at first intercourse. Table 2 displays random intercept HLM models predicting all sexual outcomes. For age at first intercourse, positive coefficients reflect higher hazards—or probabilities—of having first sex in each year. The between-family estimates for each age at father departure reflect average differences between families with father always absent, father left between birth and age 5, and father left between ages 6 and 13 relative to those with father always present, controlling for child and family level demographic differences. For age at first intercourse, families whose fathers left at any time had significantly higher hazards than those whose fathers were always present. Associations were similar in magnitude across ages at father departure. The within-family estimates were smaller in size but told a similar story. Youth whose fathers were always absent had significantly higher hazards than siblings whose fathers were always present, as did youth whose fathers left between ages 6 and 13 relative to siblings with fathers always present. No within-family association emerged, however, between having a father leave between birth and age 5 and age at first intercourse.

Age at menarche. Results from random intercept regression models predicting girls' age at menarche are also displayed in Table 2. A marginally significant between-family association emerged between having a father always absent and age at menarche of less than a third of a year. No other significant between-family associations emerged. Moreover, no significant within-family associations emerged between father departure and age at menarche. That is, sisters with different experiences of nonresident fatherhood did not reach menarche at significantly different ages.

Sexual attitudes. No associations emerged between youths' ideal age at childbirth and experience of father absence at the between-family level once family level characteristics were held constant. The within-family estimates, however, indicate that siblings who had a father always absent or who left between birth and age 5 reported younger ideal ages at childbirth, by over one half a

Table 2
Random Intercept Models Predicting Sexual Development Outcomes From Age at Father Departure

	Age at first intercourse		Age at menarche		Age at ideal childbirth		Marriage < childbirth	
	b	*se*	*b*	*se*	*b*	*se*	*b*	*se*
Between family								
Father always absent	0.47	0.11***	−0.29	0.16+	−0.61	0.31*	−0.61	0.22**
Left birth to age 5	0.45	0.09***	−0.18	0.13	0.25	0.26	−0.19	0.20
Left age 6 to 13	0.52	0.08***	0.08	0.13	0.38	0.24	−0.24	0.19
Always present (omitted)								
Within family								
Father always absent	0.30	0.11**	0.01	0.17	−0.60	0.36+	−0.45	0.26+
Left birth to age 5	0.09	0.10	0.18	0.16	−0.59	0.33+	−0.26	0.25
Left age 6 to 13	0.30	0.10**	0.01	0.15	−0.20	0.32	0.20	0.24
Always present (omitted)								
Youth is a girl	−0.40	0.04***			−0.55	0.12***	−0.14	0.09
Year of youth's birth	0.04	0.01***	0.03	0.01*	−0.08	0.02***	−0.03	0.01*
Youth is African American	0.30	0.06***	−0.35	0.10***	−0.34	0.19+	−1.34	0.14***
Youth is Hispanic	0.02	0.06***	−0.44	0.10***	0.17	0.19	−0.59	0.14***
Youth is White (omitted)								
Mother has high school degree	−0.03	0.06	0.16	0.10	0.30	0.19	0.12	0.14
Mother has some college	−0.05	0.08	0.08	0.13	0.43	0.24+	0.21	0.19
Mother has college	−0.70	0.13***	0.26	0.20	1.05	0.36***	0.93	0.39*
Mother has less than high school degree								
Mother's age at first birth	−0.05	0.01***	0.02	0.02	0.07	0.03*	0.05	0.02*
Family income (ln)	−0.08	0.05+	−0.02	0.07	0.65	0.13***	0.54	0.11***
Family has half siblings	0.11	0.06+	−0.05	0.09	0.73	0.19***	0.13	0.14
Year	1.47	0.04***						
Year-squared	−0.08	0.00***						
Constant	−5.91	0.51***	12.31	0.74*	17.20	1.45	−3.84	1.16**
Random effects								
Variance of constant	0.41	0.04	0.36	0.06	2.09	0.30	0.93	0.19
Variance of residual			1.42	0.06	14.48	0.39		
N	5,542		1,822		4,302		4,017	

Note. Age at menarche is estimated only for girls.
$^{+}p < .10$. $^{*}p < .05$. $^{**}p < .01$. $^{***}p < .001$.

year, than siblings whose fathers were always present, although the association was only marginally significant. The size of the age at father departure coefficients follows a monotonic pattern consistent with socialization theory.

A similar albeit weaker within-family pattern emerged for an ideal age at marriage younger than ideal age at childbirth. Siblings whose fathers were always absent were less likely to report an ideal age at marriage younger than ideal age at childbirth than those whose fathers were always present. The analogous between-family association was larger and statistically significant. The within and between-family coefficients for father departure between birth and age 5 were also negative, but nonsignificant.

Differences by Gender

To explore whether associations between nonresident fatherhood and sexual outcomes varied by gender, models were run separately for girls and boys. Results are reported in Table 3. Note, the sum of *n* values across models does not equal the total *N* for the full sample models because only youth with same-sex siblings could be included.

Age at first intercourse. For girls, associations between age at father departure and age at first intercourse resembled those for the full sample model, although they were somewhat larger. Again, all three between-family father departure experiences were significantly associated with higher hazards of first sex in each year. Moreover, the positive within-family association between a sibling with a father always absent versus one with father always present was twice as large for girls as in the full sample model. The within-family association between father departure after age 6 and age at first intercourse was similar in size in the girls only and full sample models, although nonsignificant in the former. For boys, by contrast, no significant associations emerge between age at father departure and age at first intercourse. The within-family coefficients for father always absent were significantly different across gender models according to a post hoc *t* test, $t = 3.12$, $p < .01$ (Gujarati, 1995).

Sexual attitudes: Ideal age at childbirth and ideal age at marriage before ideal age at childbirth. Differences by gender also emerged for sexual attitudes. For girls, a monotonic within-family pattern between age at father departure and age at ideal childbirth emerged that was consistent with socialization theory. Siblings with father always absent reported ideal ages at childbirth over 2 years younger, on average, than siblings with fathers always present; siblings whose fathers left between birth and age 5 reported ideal ages at childbirth over 1 year younger than those with fathers always present. For boys, however, father departure at any age was unassociated with ideal childbirth age. The within-family

Table 3
Random Intercept Models Predicting Sexual Development Outcomes From Age at Father Departure, by Gender

	Age at first intercourse				Ideal age at childbirth				Marriage < childbirth			
	Girls only		Boys only		Girls only		Boys only		Girls only		Boys only	
	b	*se*	*b*	*se*	*b*	*se*	*b*	*se*	*b*	*se*	*b*	*se*
Between family												
Father always absent	0.63	0.19**	0.42	0.17*	0.29	0.55	−0.55	0.51	−0.59	0.42	−1.11	0.32**
Left birth to age 5	0.59	0.16***	0.25	0.15+	0.56	0.46	0.06	0.46	−0.32	0.37	−0.13	0.31
Left age 6 to 13	0.59	0.15***	0.53	0.14***	0.24	0.43	0.46	0.42	−0.36	0.35	0.03	0.30
Always present (omitted)												
Within family												
Father always absent	0.62	0.22**	−0.32	0.21	−2.39	0.62***	0.13	0.73	−0.67	0.50	−0.08	0.46
Left birth to age 5	0.09	0.20	−0.06	0.19	−1.40	0.58*	−0.67	0.65	−0.39	0.49	−0.32	0.44
Left age 6 to 13	0.21	0.19	0.12	0.19	−0.44	0.55	0.34	0.63	0.04	0.45	0.87	0.44*
Always present (omitted)												
Year of youth's birth	0.06	0.01***	0.04	0.01***	−0.06	0.03*	−0.12	0.03***	−0.01	0.02	−0.05	0.02**
Youth is African American	0.04	0.12	0.57	0.11***	−0.50	0.34	−0.42	0.34	−1.53	0.27***	−0.97	0.23***
Youth is Hispanic	−0.10	0.12	0.17	0.11	−0.04	0.33	0.41	0.33	−0.70	0.27*	−0.42	0.24+
Youth is White (omitted)												
Mother has high school degree	−0.05	0.11	−0.01	0.11	0.14	0.34	0.44	0.33	0.11	0.27	0.15	0.21
Mother has some college	−0.32	0.16*	0.14	0.15	0.44	0.44	0.28	0.44	0.46	0.37	−0.14	0.30
Mother has college	−0.89	0.26**	−0.72	0.24**	0.59	0.64	1.43	0.68*	0.74	0.60	1.16	0.81
Mother has less than high school												
Mother's age	−0.04	0.02*	−0.05	0.02**	0.05	0.05	0.08	0.05	0.02	0.04	0.07	0.04+
Family income (ln)	−0.07	0.08	−0.11	0.08	1.02	0.24***	0.65	0.25***	0.71	0.20***	0.36	0.17*
Family has half siblings	−0.02	0.11	0.22	0.11*	0.38	0.34	1.22	0.34***	0.31	0.27	0.28	0.22
Year	1.76	0.08***	1.45	0.06***								
Year-squared	−0.09	0.01***	−0.08	0.00***								
Constant	−7.71	0.94***	−5.52	0.85***	13.34	2.61***	16.67	2.61	−4.88	2.17*	−2.88	1.84
Random effects												
Variance of constant	0.51	0.08	0.57	0.08	3.28	0.55	1.67	0.62	1.46	0.43	0.49	0.30
Variance of residual					11.55	0.60	17.12	0.85				
N	1,755		1,890		1,280		1,421		1,165		1,319	

Note. Model *N* values do not sum to full sample *N* values because gender models could only be estimated for youth with same-sex siblings.
$^{+}p < .10$. $^{*}p < .05$. $^{**}p < .01$. $^{***}p < .001$.

coefficients for father always absent were significantly different across models, $t = 2.64$, $p < .01$.

No significant within-family differences emerged between age at father departure and ideal marriage age before ideal childbirth age for girls or boys. However, the negative coefficients for father always absent and father departure between birth and age 5 were larger for girls than boys, and larger than in the full sample models. Thus, although the coefficients were nonsignificant in the girls only model, the pattern was consistent with the results for ideal age at childbirth.

Sensitivity Analyses

It is possible that including half siblings in models allowed genetic or environmental differences between half siblings to bias within-family associations reported above. To gauge this possibility, models with significant within-family associations in the full sample were rerun excluding families with half siblings (see Appendix A in the supplemental material). Because cell sizes for within-family estimates were quite small, particularly those for families with one father present and one father absent siblings, more attention is paid to coefficients' size than their statistical significance. For age at first intercourse, full sibling only coefficients were similar in size to those from the all sibling model ($b =$ 0.27 vs. 0.30 for father always absent), albeit with much larger standard errors. Post hoc tests did not reveal significant coefficient differences across models. Moreover, in the full sibling model with only girls, within-family coefficients were *larger* than in the all sibling models, although not significantly (see Appendix A in the supplemental material). However, for ideal age at first childbirth and ideal marriage age before childbirth, within-family coefficients in full sibling and full sibling girls only models were smaller and in many cases reversed in sign when compared with all sibling models.

Analyses were also run with a quantitative age at father departure measure to assess the impact of using the categorical version. Results were substantively unchanged using this specification. For example, the coefficient for age at father departure predicting first intercourse was positive and significant at the trend level, $b = -.011$, $se = .006$, $p = .054$, and weaker than the within-family categorical estimates because of the nonmonotonic association between age at father departure and age at first sex. The coefficient for age at father departure predicting age at menarche was nonsignificant and near zero, $b = -.002$, $se =$ .009. The coefficients for age at father departure predicting attitudes were significant at the $p < .05$ level and in the expected direction in the full sample, but nonsignificant when the sample was reduced to full siblings.

Discussion

Although previous research has implicated nonresident fatherhood in adolescent sexual behavior, it is unclear whether these links reflect a causal chain from nonresident fatherhood to risky sexual behavior and, if so, what mechanism accounts for the impacts. The present study used a comparison of siblings approach to address both questions. With regard to the question of causality, having a father always absent or leave home between ages 6 and 13 predicted an earlier age at first intercourse. With regard to the question of mechanism, results were most consistent with the theory that weak parental monitoring explains the link between father departure and riskier sexual behavior, although equivocal support was provided for PIT and socialization. Overall, findings offer no clear answer to the question of mechanism.

The Question of Causality

By comparing siblings discordant for nonresident fatherhood, this study reduced the influence of unobserved environmental and genetic factors that typically confound comparisons of unrelated youth. Using this conservative approach, the experience of father absence from the home, and father's departure during middle childhood, predicted earlier age at first sex. Notably, this pattern obtained even when only full siblings were compared, suggesting the father absence–sexual behavior link is not genetic in origin. The former association likely reflects the impact of nonmarital childbirth on children's sexual behavior, rather than divorce, for fathers are most likely to never live with a child if parents are unwed at the time of birth. The latter association likely reflects the impact of divorce because most cohabiting relationships either end or become marriages within 3 years of the child's birth (McLanahan & Beck, 2010). By contrast, no within-family association emerged between father departure and age at menarche, suggesting the between-family association reflects genetic or environmental confounds. Using a behavioral genetic approach, Rowe (2002) found that age at menarche was more heritable than sexual debut, although both outcomes were genetically influenced. This finding taken together with those of the present study suggests that genetic differences between families drive the link between father departure and age at menarche, whereas father departure may contribute to riskier sexual behavior.

Father departure was not associated with age at first intercourse across all ages, however. Youth whose fathers left between birth and age 5 did not have a younger age at first intercourse than siblings whose fathers were always present. These differential timing effects may account for the null within-family effects of nonresident fatherhood that Mendle et al. (2009) report, for averaging across ages at father departure could obscure the impacts of very early and later departure. Timing effects could also explain, at least in part, why Ellis et al. (2012) found no main effect of father departure on sexual behavior among sisters: their analyses excluded sisters whose fathers were always absent, the group for whom within-family effects were strongest in the present study. As for the effect of late departing fathers, it is possible higher quality fathers, as Ellis and colleagues define them, stay longer in families, thus the effect of late-leaving fathers is more likely to reflect the effect of losing a high quality father, which Ellis found promoted riskier sexual behavior among girls.

The question remains as to why youth whose fathers left the home in early childhood period would not experience the same effects as their siblings. It is possible that these youth were more likely than those whose fathers were always absent or left later on to develop a stable relationship with a stepfather, and that these relationships ameliorated the effects of father absence. Youth whose fathers were always absent may be more likely to have been born to unwed parents (see above), and mothers are less likely to remarry after a nonmarital childbirth than a marital one (Bzostek, McLanahan, & Carlson, 2012; Lundberg & Rose, 2003); if the child was between 6 and 13 when her parents divorced or separated, she may not have had enough time before adolescence to build a buffering relationship with a stepfather even if her mother remarried. Future research on nonresident fatherhood and adolescent sexual behavior should investigate the potential buffering effect of a stable relationship with a stepfather.

The Question of Mechanism

Mechanisms that might account for the link between nonresident fatherhood and earlier age at first intercourse were explored in two ways: (a) by comparing effects across age at father departure from the home and (b) comparing effects across sexual outcomes. The age pattern did not clearly support one mechanism over another. The PIT predicts that both father always absent and father departure during the first 5 years would be associated with earlier age at first intercourse, yet only the former predicted age at first intercourse. However, the association between father always absent and age at first intercourse could support the PIT. The age pattern could also support either socialization or monitoring theory if we assume a unique experience—such as the presence of a stable stepfather—distinguishes those whose fathers left between birth and age 5.

Comparing results across sexual outcomes suggests a somewhat clearer answer to the question of mechanism. First, no within-family association emerged between father absence and age at menarche for girls. Because the PIT posits earlier pubertal timing as an aspect of sexual development also influenced by father departure, one signaling an accelerated reproductive strategy, these findings do not support the PIT. Other studies too have failed to find an association between nonresident fatherhood and earlier age at menarche (Kiernan & Hobcraft, 1997; Mendle et al., 2006), although Tither and Ellis (2008) did find an association. It has been suggested that although earlier pubertal timing may be an evolutionarily adaptive response to nonresident fatherhood, the rise of other environmental risk factors, such as poor nutrition leading to higher rates of childhood obesity, have lowered the average age at menarche, making it hard to detect the unique influence of one risk factor (Belsky et al., 1991). If so, youth's evolutionary-biological response to nonresident fatherhood would impact sexual attitudes and behavior more clearly than pubertal timing. In this way, the PIT could be considered an "ultimate" level theory, one which explains why sexual attitudes and behavior might respond to nonresident fatherhood as they do, whereas socialization theory is a more "proximal" pathway linking early experience to reproductive strategy (James et al., 2012). Nonetheless, support for the PIT is dubious because no effects emerged for menarche.

The findings across sexual outcomes provide equivocal support for socialization theory. A monotonic pattern emerged between age at father departure and ideal age at childbirth such that those who lived apart from their fathers longer reported younger ideal ages at childbirth, a pattern consistent with socialization theory. However, this pattern of findings did not obtain in models that excluded half siblings. Cell sizes for within-family effects were even smaller in these models than those for age at first intercourse, so imprecision could have contributed to this attenuation. It is also possible, however, that genetic or environmental differences between half siblings drove within-family links in the full sample rather than father departure per se. It is impossible to know whether genes or environment contributed more to this possible bias because half siblings theoretically differ in both ways by father departure status. However, these findings do undermine the notion that socialization explains the link between father departure and sexual behavior.

Dubious for PIT or socialization leaves monitoring as the most plausible mechanism linking father departure to sexual behavior. For youth whose fathers leave immediately or later, the lack of monitoring and social control that can accompany single parenthood may account for their earlier age at first intercourse. This mechanism may not impact children whose father left after birth but during early childhood because their mothers may have time prior to the youth's adolescence to repartner and may be more likely, for reasons explained above, to do so than never married mothers. It is also possible, however, that the impact of early and later father departure on sexual behavior may not stem from weaker monitoring, but from a mechanism not considered initially: emotional distress of father absence or family disruption. Indeed, Cavanagh et al. (2008) found that later family disruption predicted a higher number of romantic relationships in adolescence, another indicator of risky sexual behavior, and posited this explanation. To test this pathway, future research should explore the effects of age at father departure, using a within-family design, on adolescent mental health outcomes such as depression and anxiety.

Gender Differences

The significant effects of nonresident fatherhood on age at first intercourse emerged exclusively for girls. However, the unique influence of nonresident fatherhood for girls did not appear to be biological in nature: nonresident fatherhood had no within-family effect on age at menarche. It is possible that girls' attitudes toward sex and reproduction are more vulnerable to the effects of father absence than boys'; however, effects on girls' sexual attitudes did not obtain in models excluding half siblings. Rather, girls may be more impacted than boys by weak monitoring during adolescence. Alternatively, girls may be more distressed emotionally as a result of father absence or family disruption because they are more attuned to relationships and relationship quality than boys (Amato, 1993; Crockett & Randall, 2006); they may, in turn, look to early romantic relationships to address that distress for reasons unrelated to attitudes about sexual behavior. Future research should investigate the effect of nonresident fatherhood on girls' emotional well-being using a comparison of siblings approach.

Limitations

The sibling comparison approach has several limitations one must consider before drawing theoretical implications from these findings. As mentioned above, comparing half siblings does not eliminate environmental or genetic differences that could confound associations between nonresident fatherhood and sexual outcomes, whereas comparing full siblings does. However, full siblings who differ substantially on age at father departure are rare, thus it is unclear to what extent within-family estimates based solely on that comparison are generalizable to a broader population. The disadvantaged nature of the CNLSY-YA also limits generalizability. It is also possible that the results partially reflect the impact of having a sibling with a very different experience of paternal investment and the influence of that filial comparison on sexual identity (East & Jacobson, 2000). In that way, within-family comparisons may not estimate the pure effect of father absence.

It is also important to note that the definition of nonresident fatherhood used in the present study did not consider the various levels of involvement nonresident fathers have with their children. Indeed, the term "father absence," which is used in this study, belies the high levels of investment many biological fathers make, emotionally and economically, in their noncustodial children (Cabrera, Ryan, Mitchell, Shannon, & Tamis-Lemonda, 2008). To the extent that the effect of nonresident fathers on adolescent sexual development hinges on their lower level of involvement relative to resident fathers, a difference that studies have repeatedly shown (e.g., Carlson & Corcoran, 2001), average associations may obscure variation in effects by level of father involvement. It is also possible that stable stepfathers may buffer children against the effects of early family disruption, an effect that would have been obscured in this study. The role of nonresident and stepfather involvement was beyond the scope of the present study. However, future research should explore whether the effects observed are smaller, or disappear, in families in which fathers have frequent contact or close relationships with their noncustodial children or in which stepfathers form long-term, positive relationships with nonbiological children.

Summary

This study used a comparison of siblings approach to explore the impact of nonresident fatherhood on adolescent sexual behavior. Because this approach substantially reduced the confounding effects of unobserved environmental and genetic factors that vary between families, the results strongly suggest that nonresident fatherhood, beginning either at birth or during middle childhood, leads to an earlier sexual debut for girls, but not for boys, an effect likely explained by weaker monitoring or, possibly, emotional distress. Future research should explore the buffering role that stable stepfathers, or highly involved nonresident fathers, may play in this apparent pathway. If father involvement moderates effects, it would suggest that programs to promote responsible fatherhood should encourage father involvement, not marriage or family formation per se. If stepfather presence moderates effects, it would suggest the potential benefit of encouraging social father involvement within these programs. Overall, the present study suggests all efforts to reduce the prevalence of risky sexual behavior among adolescents should consider the role nonresident fatherhood plays in its etiology.

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