COPARENTING AND NONRESIDENT FATHERS' INVOLVEMENT WITH YOUNG CHILDREN AFTER A NONMARITAL BIRTH*

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We use data from the Fragile Families and Child Wellbeing Study to investigate the association between coparenting quality and nonresident fathers' involvement with children over the first five years after a nonmarital birth. We find that about one year after a nonmarital birth, 48% of fathers are living away from their child, rising to 56% and then to 63% at three and five years, respectively. Using structural equation models to estimate cross-lagged effects, we find that positive coparenting is a strong predictor of nonresident fathers' future involvement, whereas fathers' involvement is only a weak (but significant) predictor of future coparenting quality. The positive effect of coparenting quality on fathers' involvement is robust across several techniques designed to address unobserved heterogeneity and across different strategies for handling missing data. We conclude that parents' ability to work together in rearing their common child across households helps keep nonresident fathers connected to their children and that programs aimed at improving parents' ability to communicate may have benefits for children irrespective of whether the parents' romantic relationship remains intact.

Nonmarital childbearing has increased dramatically during the past several decades, with the fraction of births occurring outside of marriage rising sixfold in the latter half of the twentieth century (Ventura and Bachrach 2000). Today, fully 37% of all births in the United States are to unmarried parents, with even higher proportions occurring among racial and ethnic minorities (Martin et al. 2006). Although many unmarried parents are cohabiting when their child is born (Graefe and Lichter 1999), about half of these parents will be living apart by their child's third birthday (Osborne and McLanahan 2007). These trends, along with the growing recognition of the potential benefits of fathers' involvement for children (Cabrera et al. 2000; Lamb 2004), have stimulated interest among researchers and policymakers in the factors that promote fathers' involvement after the romantic relationship between the parents has ended.

This paper addresses two questions: What is the prevalence of fathers' involvement with nonresident children after a nonmarital birth? And, what is the effect of a high-quality coparenting relationship on fathers' involvement? A high-quality coparenting relationship is described as one in which the parents agree about how their child should be raised, cooperate in carrying out shared objectives, and demonstrate mutual support and commitment in rearing their common child (McHale 1995). The coparenting relationship is distinct from the parent-child relationship and can exist regardless of the parents' romantic involvement or marital status (Hayden et al. 1998; Schoppe-Sullivan et al. 2004).[1]

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1. However, since the "meaning" of coparenting and fathers' involvement are qualitatively different when fathers live with—versus live away from—the mother and child, pooling resident and nonresident fathers in the

Understanding the link between coparenting and fathers' involvement after a nonmarital birth is important for several reasons. We know that children born to unwed parents are typically economically disadvantaged (DHHS 1995; Wu and Wolfe 2001); thus, fathers' contributions of time and money may represent a substantial resource for these children and their mothers (Amato and Gilbreth 1999; King and Sobolewski 2006). Also, a cooperative, low-conflict coparenting relationship may affect child outcomes directly, although to our knowledge such a link has not been demonstrated except among coresident parents (Belsky, Putnam, and Crnic 1996; Cowan and McHale 1996; McHale, Johnson, and Sinclair 1999; Schoppe, Mangelsdorf, and Frosch 2001) and divorced parents (Maccoby and Mnookin 1992). Finally, this topic has policy implications insofar as the U.S. government is currently spending money through the Healthy Marriage Initiative on programs designed to promote marriage and fathers' involvement by improving relationship skills and relationship quality among unmarried parents (Dion 2005). While such programs are targeted on romantically involved couples, the relationship skills learned through these programs may also facilitate better cooperation among couples around parenting after their romantic union ends. Experimental research on divorced parents has demonstrated long-term increases in positive coparenting and parental involvement for couples who mediate—rather than litigate—their divorce agreement (Emery et al. 2001; Maccoby and Mnookin 1992), suggesting that improving the quality of couple interactions around the time of separation may have long-term benefits for parents and children. Therefore, to the extent that relationship skills training also improves coparenting, our study sheds light on a potential ancillary benefit of marriage promotion programs: keeping nonresident fathers connected to children.

Our analysis extends previous research in several ways. First, we use a large national sample of unmarried fathers, whereas most previous research is based on smaller samples of (mostly) divorced fathers. Second, we focus on a recent birth cohort of young children and examine the coparenting relationship shortly after separation, whereas previous research has focused on adolescents and examined coparenting long after the couple relationship has ended. Finally, our analysis addresses issues of causality and feedback effects between coparenting and fathers' involvement. In sum, this paper provides important new information about family processes for a growing share of children who are born outside of marriage and who are expected to live apart from their biological fathers during early childhood.

THEORETICAL PERSPECTIVES

Family systems theory stresses the importance and dynamic nature of various family relationships (mother-father, parent-child, and sibling-sibling) that affect each other and influence individual outcomes (Bronfenbrenner 1986; Minuchin 1988). Among these dyadic relationships, an important family-level (or triadic) relationship is the one between adults who are raising a child together. The coparenting relationship is defined as the extent to which parents can effectively work together in rearing their common child and has been identified as a unique construct that is distinct from both couple relationship quality and parenting behavior (Hayden et al. 1998; McHale 1995; McHale et al. 2000). The importance of coparenting for family life is underscored by its description in a classic family text as "the family's executive subsystem" (Minuchin 1974, as cited in Schoppe-Sullivan et al. 2004). *Coparenting* has been differentiated from *parallel parenting,* in which each parent maintains a relationship with their child separate and distinct from that of the other parent (Furstenberg 1988; Furstenberg and Cherlin 1991).

For parents living apart, coparenting may represent the primary—or only—regular interaction they have with each other as they endeavor to coordinate their parental investments across households with respect to their common child (Margolin, Gordis, and John

analyses is problematic; much of the literature on fathers' involvement has looked separately at resident versus nonresident fathers, and we follow that approach here.

2001). Indeed, cooperative parenting may take on even greater import when families do not share the unifying context of household residence (Maccoby, Depner, and Mnookin 1990). In addition, both quantitative and qualitative research suggests that children's ties to nonresident fathers are tenuous over time, particularly if the father never or only briefly lived with the child (Furstenberg and Harris 1993; Lerman 1993).

We draw on several arguments that suggest how coparenting quality and fathers' involvement should be linked. These arguments hinge on the fact that mothers typically have custody of children when parents live apart, which means that fathers' access to their children is highly contingent on mothers' approval. Sometimes referred to as *gatekeeping*, mothers' facilitation of fathers' involvement can have a profound effect on fathers' roles in childrearing (Allen and Hawkins 1999), especially for parents who live apart (Ahrons and Miller 1993). Although it is important to note that mothers often facilitate rather than hinder fathers' involvement (Walker and McGraw 2000), it is clear that custodial mothers are typically able to make choices about when and how the father spends time with his child, particularly when the child is young. When parents can communicate effectively and when the mother trusts the father and believes he has the child's best interest at heart, she is more likely to encourage and support the father's active involvement because she believes his investments will be beneficial to the child. By contrast, when the mother is not able to cooperate effectively with the father and does not perceive that they are a "team" in their parental obligations, she may discourage his involvement.

Also, we have reason to expect reciprocal effects going from fathers' involvement to the coparenting relationship. As the father demonstrates greater paternal effort and spends more time with the child, the mother may develop greater trust in his intentions and greater confidence in his parenting capabilities and awareness of the child's needs, thereby facilitating effective coordination of childrearing responsibilities across households. In more general terms, the father's involvement with the child represents a "relationship-specific investment" that is unique to the biological parents of a particular focal child, enhancing social capital and positive relationships within the (non-intact) family (England and Farkas 1986). Ultimately, feedback effects might operate such that coparenting and nonresident fathers' involvement are mutually reinforcing over time.

EMPIRICAL RESEARCH

Fathers' Involvement After Nonmarital Birth

While the majority of research on nonresident fathers has focused on divorced fathers (Argys and Peters 2001), several studies in the past decade have begun to examine involvement by unmarried nonresident fathers. Using data from the National Longitudinal Study of Youth (NLSY) 1979 cohort, Lerman and Sorenson (2000) found that four years after a nonmarital child was first observed in the survey (0–2 years after the birth), 60% of fathers were living away from that child, and of these fathers, only 56% had seen their child in the past month. Using data from the NLSY 1997 cohort, Argys and Peters (2001) found that 57% of never-married nonresident fathers had any contact with their child in the past year, and 19% had weekly contact. A recent overview with data from six large data sets that included nonmarital births found that 45%–62% of white fathers and 39%–81% of nonwhite fathers had any contact in the previous year with their child age 5 or under (Argys et al. 2007).

Coparenting and Fathers' Involvement

Spurred by the rising divorce rates in the 1970s, the initial research on coparenting across households focused on parental relationships following divorce, emphasizing the deleterious effect of postdivorce conflict for children and the importance of parents' working together to rear their child even while living separately (Ahrons 1981; Wallerstein and Kelly

1980). In a major study using a representative sample of more than 1,000 postdivorce families in California, Maccoby and Mnookin (1992) identified both conflict and cooperation as two key aspects of coparenting (with some parents disengaging altogether) and concluded that children benefit from cooperative coparental relationships and are adversely affected by conflicted coparental relationships; indeed, coparenting has been shown to have both a positive and a negative dimension (Sobolewski and King 2005).

The nature of parents' divorce process matters for coparenting: as noted earlier, a randomized experiment by Emery and colleagues (2001) showed that mediation promotes longer-term coparenting by nonresidential parents than does litigation. Establishing a coparental relationship appears to be particularly important in the period immediately following divorce, setting the trajectory for whether the father stays involved (Ahrons and Miller 1993). Other research indicates that parents may avoid contact with each other in order to minimize conflict (McLanahan et al. 1994; Seltzer, McLanahan, and Hanson 1998). Although cooperative coparenting after divorce is relatively rare, when it does occur, it is shown to increase the father's role in childrearing decisions (Furstenberg and Nord 1985).

Recent research has focused on coparenting among coresident (mostly married) families and how it relates to couples' relationship quality, parenting behavior, and child well-being. This growing literature, published mostly within psychology, is based primarily on small, nonrepresentative samples. Findings underscore the distinct nature of coparenting—vis-à-vis both couple relationship quality and parenting (Hayden et al. 1998)—and suggest that among coresident households, coparenting is linked to both marital behavior (Belsky and Hsieh 1998; Schoppe-Sullivan et al. 2004) and child well-being (Schoppe et al. 2001) and may, in fact, mediate between the former and the latter (Katz and Low 2004; Margolin et al. 2001). Cooperative coparenting within two-parent families has been linked to more responsive parenting by both mothers and fathers with infants and school-age children (Caldera and Lindsey 2006; Floyd, Gilliom, and Costigan 1998; Margolin et al. 2001).

To our knowledge, only one study has examined coparenting and nonresident fathers' involvement using a nationally representative data set. Sobolewski and King (2005) used data from the National Survey of Families and Households to determine how coparenting and conflict over childrearing affect father-child contact, the quality of father-child relationships, and responsive fathering among nonresident fathers. Their sample included parents of children and adolescents aged 10–18 who were born between 1974 and 1984, and the vast majority (84%) of parents were divorced (as opposed to never married). They found that cooperative coparenting (but not conflict) was positively linked to all three measures of fathers' involvement, with father-child contact mediating the effect of coparenting on the quality of the father-child relationship and responsive fathering.

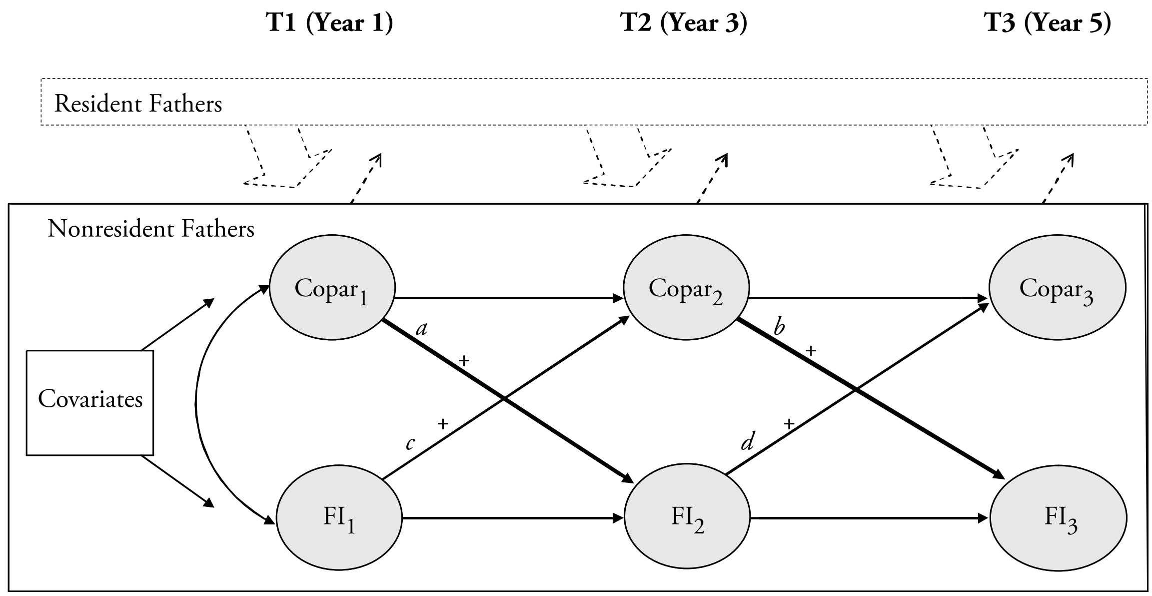
THE PRESENT STUDY

Conceptual Model

Figure 1 presents a diagram of our conceptual model for how coparenting and fathers' involvement are linked over five years after a nonmarital birth. The figure indicates that among the population of nonresident fathers, the coparenting relationship and fathers' involvement have reciprocal effects: coparenting at a point in time affects fathers' future involvement (pathways *a* and *b*), and fathers' involvement at a point in time affects future coparenting (pathways *c* and *d*). The figure also indicates that there are flows into and out of residence status, with some fathers moving from residence to nonresidence at one, three, and five years, and others (a smaller group) moving from nonresidence to residence.

The conceptual model highlights several challenges that must be addressed in the empirical analyses in order to obtain an unbiased estimate of the effect of coparenting on nonresident fathers' involvement. First, it highlights the fact that the association between

Figure 1. Conceptual Model for Residence Status, Coparenting, and Father Involvement



Note: Copar = coparenting. FI = father involvement.

coparenting and fathers' involvement is likely bi-directional: coparenting quality is expected to increase fathers' involvement, and fathers' involvement is expected to increase coparenting quality. Based on past research, we expect a stronger link between coparenting and involvement than vice versa. However, if the effect of fathers' involvement on coparenting is strong and positive, this will lead to an upward bias in our estimate of the effect of coparenting on fathers' involvement if we do not explicitly estimate the reciprocal pathways. To address this problem, we use structural equation models to examine the cross-lagged associations between coparenting and fathers' involvement at three time points.

Second, the model highlights the fact that there is selection into and out of the status of "nonresident father" over time. Insofar as unmeasured variables affect fathers' residence status as well as the quality of the coparenting relationship and fathers' involvement, conditional on nonresidence, these variables must be taken into account. We use several strategies to deal with possible selection bias. First, we estimate standard ordinary least squares (OLS) models, using a rich set of observed covariates to directly control for factors that are likely to affect fathers' residence status as well as the other variables of interest. Next, we estimate a Heckman two-stage model designed to correct for selection bias due to both observed and unobserved heterogeneity (Heckman 1979), likely reducing the magnitude of the coefficients. Finally, along with random-effects models that reflect both between- and within-person differences, we estimate fixed-effects models that allow us to control for unobserved differences among nonresident fathers that are constant over time.

Moderating Factors

Coparenting may be more or less strongly predictive of fathers' involvement for particular subgroups, so we explore several variables that may moderate the association. First, we examine child gender because there is some evidence that fathers are more involved with sons than with daughters, especially at older child ages (Cooksey and Fondell 1996; Harris and Morgan 1991)—although other studies find no such differences (Cooksey and Craig

1998; Seltzer 1991). To the extent that fathers may identify more strongly with same-sex children (Rossi and Rossi 1990), we expect that coparenting may be more important for keeping fathers involved with girls.

Second, there is reason to believe that coparenting and fathering processes may operate differently by fathers' race/ethnicity. Being a nonresident father is more normative among African Americans, with 69% of all births occurring outside of marriage (Martin et al. 2006). Some studies have shown that black nonresident fathers are more involved with their children compared with other racial/ethnic groups (King 1994; Seltzer 1991); compared with white fathers, black fathers' residential patterns are more fluid, and their involvement is more likely to persist in the face of an ambiguous relationship with the child's mother (Mott 1990). Nonresident fathers' involvement may be less contingent on coparenting for black fathers than for other fathers.

Third, nonresident fathers' involvement with a given child is affected by the presence of children from his other partnerships (Manning and Smock 1999; Manning, Stewart, and Smock 2003). To the extent that previous children (of fathers and perhaps mothers) may diminish fathers' ability to both coparent with mothers and invest in their common child, we expect that the link between coparenting and involvement may be weaker when fathers have one or more previous children.

Control Variables

In order to minimize spurious associations, we control for a number of factors that may be linked with nonresident fathers' involvement and are likely to affect parents' ability to coparent effectively. Compared with resident fathers, nonresident fathers are typically younger, less educated, in worse physical health, more likely to be depressed, and more likely to abuse substances (DeKlyen et al. 2006; Garfinkel, McLanahan, and Hanson 1998). Nonresident fathers' having access to children—and the extent of involvement when they do—is shown to be a function of mothers' perceptions about the father as a person and parent (via *gatekeeping*) and a host of father and child characteristics (Cooksey and Craig 1998). Key paternal characteristics include age (Landale and Oropesa 2001), socioeconomic status (Cooksey and Craig 1998; Seltzer and Bianchi 1988), family background (McLanahan and Sandefur 1994), religiosity (King 2003; Wilcox 2002), and obligations to new children and partners (Manning and Smock 1999; Seltzer and Bianchi 1988), as well as physical health, mental health, social-behavioral problems (such as having a substance abuse problem and being physically violent), and having a history of incarceration (Western, Lopoo, and McLanahan 2004). Race/ethnicity is related to nonresident fathers' involvement, though not always consistently (King, Harris, and Heard 2004). With respect to child characteristics, as noted above, there is mixed evidence about how child gender affects fathers' involvement (Cooksey and Craig 1998; Cooksey and Fondell 1996; Harris and Morgan 1991; Seltzer 1991), and a child's having a more "difficult" temperament may deter positive parenting (Simons et al. 1990).

METHODS AND DATA

Data

We use data from the Fragile Families and Child Wellbeing Study, a national longitudinal study designed to examine the characteristics of unmarried parents, the relationships between them, and the consequences for children. The study follows a birth cohort of 3,712 children born to unmarried parents (and a comparison group of married parents) in 20 U.S. cities with populations of 200,000 or more. Baseline interviews with mothers and fathers were conducted shortly after their child's birth between 1998 and 2000. Mothers were interviewed in person in the hospital within 48 hours of the birth, and fathers were interviewed in person or by phone as soon as possible thereafter, either in the hospital or wherever they could be located (Reichman et al. 2001). Follow-up interviews with both mothers and fathers were conducted when the child was about 1, 3, and 5 years old.

A completed mother interview at baseline was required for the mother/father/child to be included in the panel, but beyond this basic criterion, both parents were eligible for each follow-up survey, even if they were not interviewed in the previous wave(s). At the baseline survey, 87% of eligible unmarried mothers agreed to participate in the study, and 75% of the fathers of their child were interviewed.[2] At the one-year follow-up, 90% of eligible unmarried mothers and 70% of eligible unmarried fathers were interviewed, where eligibility is defined as a completed mother baseline interview; the response rates for the subsequent three- and five-year waves were, respectively, 88% and 87% of eligible unmarried mothers and 68% and 66% of eligible unmarried fathers.

Sample

Our analyses are based on several different subsamples of children who were born outside of marriage, who lived with their biological mother, and who had a nonresident father at some point within five years after birth. For one set of analyses (structural equation, OLS, and Heckman models), we use the 1,228 fathers who were nonresident at the one-year survey, whose child lived with the mother, and who had seen their child at least once since birth. From the 3,712 cases at baseline, 418 cases were excluded because the mother was not interviewed at the one-year survey, 52 were excluded because the child did not live with the mother at the one-year survey, and 8 cases were excluded because data were missing on father coresidence status, yielding 3,234 cases;[3] of these, only 1,592 of the fathers were nonresident at the one-year survey. An additional 210 cases were excluded because the father had not seen the child since birth, and 154 cases were dropped because the coparenting questions were not asked at the one-year survey in two cities (Austin and Oakland), yielding a final one-year sample of 1,228 cases in which the mother was interviewed (of these, 752 fathers were interviewed); listwise deletion of missing data yielded subsamples of 896 complete cases on mother-reported variables and 690 complete cases on both mother- and father-reported variables. For the outcomes of spending time and engaging in activities, we focus on the subset of 875 fathers who saw their child more than once in the previous month (again, with slightly smaller subsamples when we use listwise deletion to get complete cases for each outcome); these are the fathers for whom measuring variation in fathering is salient.

The subsamples of cases do not notably differ from the full sample of nonresident fathers on most characteristics included in the analyses. The full sample includes a slightly higher share of black fathers, fathers who have a child by another partner, and fathers who have been previously incarcerated; for example, 50% of the 1,592 nonresident fathers have a history of incarceration, compared with 45% of the 690 complete cases. The biggest difference across the samples is that a smaller proportion of the full sample was in a romantic relationship and/or cohabiting with the mother at the time of birth compared with the subsamples; for example, 25% of the full sample (1,592) were cohabiting at birth, and 44% were in a romantic noncohabiting relationship (68% total), compared with 36% cohabiting and 50% in a romantic noncohabiting relationship (86% total) among the smallest subsample (690) that is limited to father-interviewed cases. This is not surprising, given that we are limiting all analyses to cases in which the father had some contact with the child since birth, and also, the father-interviewed cases are most closely connected to the mother (and child).

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2. The Fragile Families data are the most representative of cohabiting fathers (90% response rate) and the least representative of fathers who were not romantically involved with the child's mother at the time of birth (38% response rate).

3. Of these 3,234 cases, 1,678 (or 52%) were nonresident at the baseline survey.

For our random-effects and fixed-effects models (see below), we pool cases across the one-, three-, and five-year surveys when the mother was interviewed, the child lived with the mother, and the father was nonresident but had some contact with the child since the previous interview, for a total of 4,062 person-year observations, or 2,191 unique cases (of these, 1,607 fathers were interviewed); the mean number of survey waves for which each case is observed is 1.85. Listwise deletion yields 1,657 complete cases on mother-reported variables and 1,110 on both mother- and father-reported variables. For the outcomes of spending time and engaging in activities, we focus on the subset of 2,866 person-year observations (1,733 unique cases) in which the father saw the child more than once in the previous month; again, listwise deletion of missing data yields a smaller number of complete cases for each outcome.

Variables

We measure cooperative coparenting between mothers and fathers by using a series of six items about how the parents work together in raising their child, reported by mothers at each of the one-, three-, and five-year interviews if the father saw the child at least once since the previous survey. These items are (1) “When (father) is with (child), he acts like the father you want for your child,” (2) “You can trust (father) to take good care of (child),” (3) “He respects the schedules and rules you make for (child),” (4) “He supports you in the way *you* want to raise (child),” (5) “You and (father) talk about problems that come up with raising (child),” and (6) “You can count on (father) for help when you need someone to look after (child) for a few hours.” Response choices are “rarely true” (1), “sometimes true” (2), and “always true” (3).[4] Factor analysis confirmed that the items loaded on a single factor and could be appropriately averaged into an index (Cronbach’s alpha = .87 at one year, .89 at three years, and .88 at five years).

We measure fathers’ involvement by using several identical measures from the one-, three-, and five-year surveys, all reported by mothers. We use mothers’ reports of fathers’ involvement in order to examine a greater share of fathers, since mothers are more likely to be interviewed than nonresident fathers (we discuss below how we deal with missing data). We present descriptive information about whether the father saw the child since the previous wave and whether the father saw the child more than once in the previous month (yes/no). Our first measure for analysis is the number of days the father saw the child in the past month, ranging from 0 to 30. Our second measure, reported if the father saw the child more than once in the past month, indicates how often the father spent one or more hours with the child in the past month, ranging from 1 (never) to 5 (every day). The third measure, also reported if the father saw the child more than once in the past month, indicates the mean number of days in the past week (0 to 7) that the father engaged in four activities with the child—singing, reading stories, telling stories, and playing with toys (alpha = .87 at one year and .90 at both three and five years).

We include a range of control variables in order to avoid spurious relationships between coparenting and fathers’ involvement. These variables measure fathers’ baseline demographic characteristics, health and human capital, and sociobehavioral characteristics, as well as several child characteristics reported by the mother or father (as indicated) at the time of the baby’s birth unless otherwise noted. In addition, we include several time-varying characteristics.

Fathers’ age at the time of birth is measured in years, reported by fathers (using mothers’ report if fathers’ is missing). Fathers’ race/ethnicity is specified as non-Hispanic black (reference), non-Hispanic white, Hispanic, and non-Hispanic “other” race (again, using fathers’ report unless missing). Immigrant status is measured by a dummy variable for

4. At the three- and five-year surveys, an additional choice of “never” was given; we combine the small number of responses in this category with “rarely true” to yield a consistent three-point scale across all years.

whether the father reports that he was born outside the United States. Family background is represented by a dichotomy for whether the father reports that he lived with both of his parents at age 15. We include a dummy variable indicating whether the father had a child by a previous partner, a variable for whether the mother had a child by a previous partner, and a continuous variable for the number of biological children the focal parents have together (all reported by mothers at the one-year survey).

Fathers' education is specified as less than high school (reference), high school diploma, and some college or above (reported by fathers if available, otherwise by mothers). Fathers report their physical health status in categories of poor, fair, good, very good, or excellent. We include a measure of fathers' self-reported attitudes toward fathering based on three items, with responses ranging from 1 (strongly disagree) to 4 (strongly agree): (1) "Being a father and raising children is one of the most fulfilling experiences a man can have," (2) "I want people to know that I have a new child," and (3) "Not being a part of my child's life would be one of the worst things that could happen to me" (alpha = .74). Fathers' religious attendance is self-reported as how often he attends religious services, ranging from 1 (never) to 5 (once a week or more).

Mothers report whether fathers have a substance problem by responding (yes/no) to the question, "Does (baby's father) have problems such as keeping a job or getting along with family and friends because of alcohol or drug use?" Physical partner violence toward the mother is represented by a dummy variable coded as 1 if the mother reported at the one-year survey that she was ever "seriously hurt" by the father at some point before the baby's birth. We include fathers' incarceration history as a dummy variable indicating whether the father has ever been in jail or prison (based on both mothers' and fathers' reports).

We include two pieces of information about the child: gender, and temperament based on three items from the Emotionality, Activity, Sociability, and Impulsivity (EASI) scale reported by mothers at the one-year survey. For the EASI items, the mother indicates whether certain statements reflect her child's behavior, ranging from 1 "not at all like my child" to 5 "very much like my child." We include the mean of three items that indicate "difficult" temperament: "he/she often fusses or cries," "he/she gets easily upset," and "he/she reacts strongly when upset" (alpha = .61); maternal reports on this measure are shown to correspond with observations from independent interviewers obtained during an in-home visit (Meadows, McLanahan, and Brooks-Gunn 2007).

We also include several time-varying covariates (measured at one, three, and five years) in the pooled analysis: fathers' self-reported work hours (measured as the total number of hours worked at all jobs in the past week) and whether the father or mother each reports having a new romantic partner. Since these variables may be affected by fathers' involvement as well as the coparenting relationship, including them in our model makes our estimates of the effects of coparenting on fathers' involvement more conservative.

Table 1 provides descriptive information about our primary (pooled) analytic sample of nonresident fathers subsequent to a nonmarital birth (2,191 fathers). The average unmarried father was in his mid-20s when his baby was born and is of minority race/ethnicity: 52% of fathers are non-Hispanic black, and 32% are Hispanic. Only 10% are immigrants. Thirty-seven percent lived with both parents at age 15. Three-fourths have a high school diploma or less. On average, these fathers have 1.4 children with the biological mother. Over two-fifths of these fathers had a child by another partner at the time of the focal child's birth, and 37% of mothers had a previous child. Most fathers hold positive attitudes toward being a father and attend church infrequently. Most fathers are in good health, do not report a substance problem, and are not physically violent. Yet, fully 44% of fathers have been previously in jail/prison. About half the children are boys, and the average child temperament falls in the high middle range of the "difficult" scale.

With respect to the variables measured over time, the average level of coparenting at Year 1 is 2.3 (on a 1-to-3 scale), declining to 2.1 at years 3 and 5. Nonresident fathers work

Table 1. Characteristics of Nonresident Fathers

Variables	% or Mean	SD
Baseline Characteristics		
Father's age at baby's birth (mean)	25.75	7.78
Father's race/ethnicity		
White non-Hispanic	11.3	
Black non-Hispanic	52.2	
Hispanic	31.5	
Other non-Hispanic	5.0	
Father immigrant (born outside the U.S.)	10.4	
Father lived with both parents age 15	36.8	
Father's education		
Less than high school	35.5	
High school	41.7	
Some college or more	22.8	
Number of biological children with mother (mean)	1.35	0.72
Father has prior child(ren) by other partner	44.2	
Mother has prior child(ren) by other partner	36.5	
Pro-fathering attitudes (mean; range = 1–4)	3.63	0.51
Father's religious attendance (mean; range = 1–5)	2.66	1.30
Father's self-reported health (mean; range = 1–5)	4.00	0.98
Father has substance problem	7.7	
Father is physically violent toward mother	5.6	
Father has history of incarceration	43.6	
Baby is a boy	51.5	
Child's "difficult" temperament (mean, range = 1–5)	2.86	1.08
Number of (unweighted) person-year observations	4,062	
Number of (unweighted) unique cases	2,191	

	Year 1		Year 3		Year 5	
Time-Varying Characteristics	% or Mean	SD	% or Mean	SD	% or Mean	SD
Coparenting (mean; range = 1–3)	2.30[a]	0.66	2.08	0.69	2.12	0.67
Father's work hours per week (mean)	33.19	25.74	30.58	25.13	35.54	25.38
Father has a new (current) romantic partner	30.4		53.0		49.5	
Mother has a new (current) romantic partner	24.8		37.4		45.6	

Notes: All figures are weighted by national sampling weights. Data are pooled across the one-, three- and five-year surveys and include all fathers who were unmarried at the time of the focal birth (with valid mother interviews), were nonresident at one or more survey waves, and saw the child at least once since the previous survey.

[a]Values for Austin and Oakland are not included because these questions were not asked in these cities at the one-year survey.

an average of 30 to 36 hours per week. At Year 1, 30% of fathers (and 25% of mothers) have a new partner, rising to 50% of fathers and 46% of mothers by Year 5.

Missing Data

Missing data are of concern in any observational study, and the proper treatment of missing values is important for obtaining unbiased estimates (Allison 2002). In this paper, we present regression estimates using three approaches to handling missing data: first, we use listwise deletion to obtain complete cases on all father- *and* mother-reported variables; next, we use listwise deletion to obtain complete cases on all mother-reported variables (which allows us to include cases with noninterviewed fathers); and third, we use multiple imputation (MI) to fill in missing values for our full sample of nonresident fathers. Introduced by Donald Rubin in the 1970s, MI has emerged as a promising strategy for dealing with missing data that eliminates the biases inherent in more conventional methods, such as mean substitution or dummy variable adjustment (Allison 2002; Rubin 1976, 1987). MI uses observed data to impute missing values over multiple data sets; analyses are then conducted across each data set, and the estimates are averaged to reflect the intrinsic uncertainty in the missing-data imputation (and hence yield appropriate standard errors).

There are only two mother-reported variables in our data with more than 10% of cases missing before imputation—father's having children by another partner (14%) and father's violence (18%). For the father-reported variables, about 30% of cases are missing from the total sample, reflecting the lower response rate for fathers compared with mothers. We used MI techniques to generate estimates for the missing values; in the imputation model, we include variables reported by mothers and fathers that are (a) related to the substantive question within this research (how coparenting affects fathers' involvement) and/or (b) related to the likelihood of being missing (Allison 2002). Therefore, we include the variables for coparenting, our father involvement measures, control variables associated with both, as well as variables associated with nonresponse (e.g., we know that couple relationship status at baseline is associated with the father's not being interviewed). We use Stata 9.2 SE with the *ice* (imputation by chained equations) command developed by Patrick Royston (Royston 2004). In our structural equation models, we handle missing data using full information maximum likelihood (FIML) within the Mplus statistical software (Muthén and Muthén 2006); FIML estimates models that include all cases using all available data and has been shown to yield less biased and more efficient estimates than other missing data treatments, such as listwise deletion and mean imputation (Wothke 2000).

Analytic Techniques

We use several analytic techniques to deal with the issues of possible feedback effects and selection on unobserved variables. First, in order to evaluate whether the quality of the coparenting relationship affects fathers' involvement, net of any feedback effects from involvement to coparenting, we estimate a cross-lagged structural equation model (using Mplus software, Version 4) with identical measures of coparenting and fathers' involvement drawn from the one-, three-, and five-year surveys. A cross-lagged design allows us to estimate the extent to which coparenting affects future fathers' involvement over and above its effect on later coparenting and net of the individual control variables, and the extent to which fathers' involvement affects future coparenting over and above its effect on later fathers' involvement and net of the control variables.

After determining the direction of the effects, we then evaluate whether the effect of coparenting on involvement is likely to be causal. Here we estimate a series of linear regression models designed to deal with possible selection associated with the status of nonresident father (and then having any contact with the child). Our first regression model uses OLS to regress fathers' involvement on coparenting quality with control variables. The Fragile Families data contain a rich set of covariates that are not typically measured in other studies, and these covariates allow us to directly control for many of the variables that are likely to affect fathers' residence as well as the coparenting relationship and levels of involvement. The OLS models provide a baseline estimate of the overall association between coparenting and nonresident fathers' involvement at the one-year survey, controlling for observed characteristics.

We next use a Heckman selection model to estimate the effect of coparenting on fathers' involvement (Heckman 1979). Here, a two-equation model is used in which the first-stage "selection equation" predicts fathers' residence, and the second-stage equation predicts fathers' involvement. For the equation predicting the number of days the father saw the child in the past month, the outcome variable in the first-stage equation is whether the father is nonresident. For the equations predicting the level of fathers' spending time and engagement with the child, the outcome variable in the first-stage equation is whether the father is nonresident *and* sees the child more than once a month. We could not think of a good instrument for the first-stage equation, and so the selection models are identified by functional form; this limitation is discussed below.

Finally, we estimate random-effects and fixed-effects models. These models take advantage of the longitudinal design of the data, using repeated observations pooled over time. The random-effects models allow us to examine the relationship between coparenting and fathers' involvement over a longer time span and with a larger sample of nonresident fathers, capturing variation both between and within subjects. The fixed-effects models utilize only within-subject variation and reflect how changes in the coparenting relationship are associated with changes in fathers' involvement with their nonresident children. This more conservative technique reduces bias in the estimates by controlling for unobserved individual characteristics that do not change over time and that may be associated with coparenting and fathers' involvement (Greene 2003; Snijders 2005); yet, these models do not address the issue of reciprocal effects between coparenting and fathers' involvement. As with the OLS and Heckman models, we estimate models for how coparenting predicts the frequency of contact and the frequency of fathers' spending time and engaging in activities, conditional on contact.

For all four of our regression techniques (with each of our three fathers' involvement outcomes), we estimate three sets of models using a slightly different sample each time, depending on the treatment of missing data. First, we limit the sample to complete cases with valid (nonmissing) data on all covariates, including those reported by mothers and fathers. Second, we limit the sample to complete cases with valid (nonmissing) data on all mother-reported covariates. Dropping the restriction on father-reported covariates allows us to increase the sample size and include cases in which the father was not interviewed. Third, we use MI techniques (described above) to impute missing data on all covariates and hence include the entire sample of nonresident fathers.

PREVALENCE OF FATHERS' INVOLVEMENT

Our first research question asks about the prevalence of nonresident fathers' involvement during the first five years after a nonmarital birth; we know that at the time of birth, 49% of unwed fathers were nonresident (McLanahan et al. 2003). As shown in Table 2, about one year after a birth, 48% of unmarried fathers were living away from the baby and mother, and the fraction rises steadily in the subsequent years, to 56% around the child's third birthday and to 63% around the child's fifth birthday. In other words, among the large (and growing) fraction of all children born outside of marriage today, more than three-fifths will be living apart from their biological father by age 5.

Among nonresident fathers, the majority maintained at least some contact with their child. At Year 1, 87% of nonresident fathers had seen their child at some time since the baby's birth, and 63% had seen their child more than once in the past month. By Year 3, 71% of fathers had seen the child since the preceding interview (around child's age 1), and

Table 2. Prevalence of Fathers' Involvement After Nonmarital Birth, Based on Mothers' Reports

	Year 1 (n = 3,234)		Year 3 (n = 3,113)		Year 5 (n = 3,037)	
	% or Mean	*SD*	% or Mean	*SD*	% or Mean	*SD*
Nonresident Fathers (%)	47.6		55.9		62.9	
All Nonresident Fathers						
Saw child since previous survey (%)	87.0		70.9		63.2	
Saw child more than once in past month (%)	62.7		47.0		43.1	
Mean number of days father saw child (range = 0–30)	8.36	10.92	6.28	9.96	5.26	9.17
Fathers Who Saw Child More Than Once in the Past Month						
Mean number of days father saw child (range = 1–30)	13.33	11.13	13.35	10.79	12.21	10.50
Mean frequency of spending one or more hours (range = 1–5)	3.70	1.24	3.67	1.14	3.54	1.06
Mean engagement in activities (range = 0–7 days)	2.08	1.79	2.10	1.70	1.51	1.80

Notes: All figures weighted by national sampling weights for each respective year. Unweighted numbers of cases (n) indicate mothers interviewed at each survey wave living with the focal child that had nonmissing data on father coresidence status.

47% had seen the child more than once in the past month. At five years, 63% of fathers had seen their child since the three-year interview, and 43% had seen the child more than once in the past month. Taken together, these figures suggest notable divergence in the level of nonresident fathers' involvement that children experience by age 5—nearly two-fifths of children (37%) had no contact with their father in the prior 1–2 years, another two-fifths (43%) had regular ongoing contact, with the remaining fifth (20%) falling somewhere in between. Among all nonresident fathers (including those who did not see the child in the previous month), the mean number of days that fathers saw their child was over 8 days in the past month at Year 1, falling to just over 5 days at Year 5.

Turning to the subset of fathers who saw their child more than once in the previous month, we find more frequent father-child contact, as expected. These fathers saw their child an average of 13 days at Years 1 and 3, and 12 days at Year 5. The frequency of spending one or more hours is close to "a few times a week" at Year 1, declining slightly over Years 3 and 5. The average number of days per week that these fathers engaged in activities with the child is 2.1 at Years 1 and 3 and 1.5 at Year 5.

We also examined differences in levels of fathers' involvement by race/ethnicity (data not shown in the table). We found that black non-Hispanic fathers were much more likely to be nonresident at each survey wave, compared with white or Hispanic fathers. Yet, among nonresident fathers, black non-Hispanic men were more likely to have maintained contact with their child, to have seen their child in the past month, and to have seen their child a greater number of days. Racial/ethnic differences on the other measures are less consistent across measures and over time.

MULTIVARIATE RESULTS FOR COPARENTING AND FATHERS' INVOLVEMENT

The primary aim of our research is to examine whether the quality of the coparenting relationship affects nonresident fathers' involvement after a nonmarital birth. Table 3

Table 3. Estimates of Reciprocal Effects of Coparenting and Nonresident Father Involvement

	$Copar_1$ to FI_3	FI_1 to $Copar_3$	$Copar_3$ to FI_5	FI_3 to $Copar_5$	χ^2	*df*	CFI	RMSEA
All Nonresident Fathers Who Saw Child Since Baby's Birth (*n* = 1,228)								
Number of days saw child (including 0s)	4.01**	0.01**	2.38**	0.00*	1,152.89	512	.943	.032
Fathers Who Saw Child More Than Once in Past Month (*n* = 875)								
Spent one or more hours with child (range = 1–5)	0.36**	0.02	0.61**	0.01	907.87	512	.945	.030
Engagement in activities (range = 1–7 days)	0.24	0.03*	0.54**	0.01	791.48	512	.961	.025

Notes: Copar = coparenting; FI = father involvement; numbers 1, 3, and 5 represent survey year; CFI = Comparative Fit Index; RMSEA = Root Mean Square Error of Approximation. Estimates are derived from structural equation models that control for the baseline characteristics of father's age, race/ethnicity, education, immigrant status, health, lived with both parents at age 15, ever incarcerated, violence, substance problem, attitudes toward fathering, number of children with mother at baseline, had prior child(ren) by other partner, as well as mother had prior child(ren) by other partner, child sex, and child "difficult" temperament. Full information maximum likelihood is used to treat missing data.

$^{\dagger}p < .10$; $^{*}p < .05$; $^{**}p < .01$

shows results for the cross-lagged paths from the structural equation models for the three outcomes of interest using identical measures of coparenting and fathers' involvement reported at one, three, and five years following the child's birth.[5] The results show a quite consistent pattern: the primary direction of the association operates from coparenting to fathers' involvement, with the effect from fathers' involvement to coparenting being much weaker. Among nonresident fathers at Year 1, each "unit" of coparenting at Year 1 predicts the father seeing the child 4.0 days more per month at Year 3, and each "unit" of coparenting at Year 3 predicts the father seeing the child 2.4 days more per month at Year 5; these represent effect sizes of about .40 and .26, respectively.[6] These estimates are net of the autocorrelations of the coparenting and fathers' involvement measures among couples across time and hence represent a rather conservative test. By contrast, fathers' involvement at Year 1 or Year 3 does not appear to have a large effect on coparenting at Year 3 or Year 5; the coefficients are both statistically significant but close to zero in magnitude.

Turning to our subsample of fathers who had contact with their child more than once in the past month, we find that coparenting significantly predicts fathers' spending one or more hours per day with their child at both Year 3 and Year 5 and significantly predicts more frequent engagement in activities at Year 5. Each "unit" of coparenting predicts a higher score on the measure for spending one or more hours per week of .36 at Year 3 and .61 at Year 5, corresponding to effect sizes of .32 and .58, respectively. Each "unit" of coparenting is associated with a higher score on engagement in activities of .24 at Year 3, although this estimate is not statistically significant (the standard error is very large). Between Years 3 and 5, each "unit" of coparenting is linked to a .54 higher score on the father-child activities measure, or an effect size of about .30. With respect to effects from fathers' involvement to coparenting, among nonresident fathers with regular contact, the magnitude of the

5. Missing data are included using full information maximum likelihood; we do not show results using only complete cases because the number of cases with no missing data across Waves 1, 3, and 5 becomes very small.

6. An effect size is the coefficient divided by the standard deviation on the mean outcome for the entire sample, often used as an indicator of the magnitude of effects. A standard typology used in behavioral research is that an effect size of .2 or less is considered small, around .5 is moderate, and .8 or above is large (Cohen 1977).

estimates is mostly small, and only one is statistically significant: each higher score on the engagement measure at Year 1 is linked with a .03 higher score on the coparenting measure at Year 3 (effect size of .04), suggesting that between Years 1 and 3 after a nonmarital birth, the father's more frequent engagement in activities with the child appears to have a very modest effect on effective coparenting between the parents. Taken together, our structural equation model results suggest that the direction of the association between coparenting and father involvement operates primarily from coparenting to involvement.

Next, we evaluate the extent to which this association can be appropriately considered causal. Here we estimate a series of OLS, Heckman, random-effects, and fixed-effects regression models using complete cases with all variables, complete cases with only mother-reported variables, and all cases using multiple imputation to impute missing values. The results, as reported in Table 4, are strikingly consistent. Looking first at the OLS estimates for the effects of coparenting on fathers' involvement, we find that each "unit" of coparenting is associated with an increase of nearly 7 days of father-child contact per month. Recall that the OLS model controls for a large set of observed covariates about fathers.

Looking at the estimates based on the Heckman model—which corrects for selection into nonresidence—the coefficient is about one-third smaller in magnitude (4.4) and just outside of the range of marginal statistical significance (p = .104). The decline between the two estimates suggests that part of the "effect" of coparenting is attributable to unobserved characteristics associated with becoming a nonresident father. However, the inverse Mills' ratio (lambda) is not statistically significant (p = .361).

Looking at the estimates based on the random- and fixed-effects models, we find that each unit increase in coparenting quality is associated with nearly 8 more days of contact per month using the random-effects model, and with 6.7 more days of contact using the fixed-effects model. Again, the difference between the two estimates indicates that some of the effect of coparenting quality on fathers' involvement is due to unobserved differences between fathers. Yet, the fixed-effects estimate remains strong and statistically significant: nonresident fathers' ability to cooperatively coparent with custodial mothers has a significant positive association with fathers' frequency of contact with their common child.

In both the random-effects and fixed-effects models, we can control for a set of time-varying characteristics that may also affect fathers' involvement and hence obtain a more conservative estimate of the effect of coparenting over time (Model 2). When we control for fathers' work hours (which may both reflect and affect fathers' perception of themselves as a breadwinners as well as their time available for parenting) and whether parents have new partners over time (which may complicate coparenting arrangements and diminish fathers' ability and motivation to see the child), we find that the association of coparenting with fathers' involvement declines by only a small amount (13%–14%).

The second set of findings in Table 4 shows estimates for the father spending one or more hours with the child among the subset of fathers who saw the child more than once in the previous month. Again, we find relatively consistent results across all regression methods. In the OLS model, each level of coparenting is linked with a score that is .87 higher on the measure of spending one or more hours (range is 1 to 5), or a relatively large effect size of .70. For this outcome, the magnitude of the effect increases going from OLS to Heckman results (instead of decreasing as with number of days) to 1.23, suggesting that the true effect of coparenting on fathers' involvement is even stronger after we account for the characteristics that predict seeing the child; the inverse Mills' ratio, however, is always outside of conventional significance levels (i.e., the p value is never less than .10). As noted above, our selection equation does not include an instrumental variable and is identified by functional form. Thus, we would not place as much weight on results from this model as we would if we had a good instrument. The fact that these results are consistent with those from the fixed-effects model (see below), however, increases our confidence that selection is not a serious problem for our analyses.

Table 4. Estimates of Coparenting on Nonresident Father Involvement

	Complete Cases, All Covariates				Complete Cases, Mother-Reported Covariates				Multiple Imputation, All Covariates			
	One-Year Data		Pooled One, Three, and Five Years		One-Year Data		Pooled One, Three, and Five Years		One-Year Data		Pooled One, Three, and 5 Years	
	OLS (a)	Heckman (b)	RE (c)	FE (d)	OLS (e)	Heckman (f)	RE (g)	FE (h)	OLS (i)	Heckman (j)	RE (k)	FE (l)
Number of Days Saw Child in Past Month (including 0s)												
Model 1: Baseline characteristics	6.76**	4.39	7.72**	6.71**	6.99**	6.12**	7.47**	6.17**	5.30**	5.01**	6.20**	4.56**
Model 2: Add time-varying variables			6.62**	5.83**							5.65**	4.09**
n (unique cases)	690	690	1,110	1,110	896	896	1,657	1,657	1,228	1,228	2,191	2,191
Spent One or More Hours (if saw child more than once in past month)												
Model 1: Baseline characteristics	0.87**	1.23**	0.94**	0.78**	0.89**	1.34**	0.98**	0.93**	0.49**	0.91*	0.63**	0.51**
Model 2: Add time-varying variables			0.85**	0.70**							0.59**	0.46**
n (unique cases)	514	514	898	898	655	655	1,315	1,315	875	875	1,733	1,733
Engagement in Activities (if saw child more than once in past month)												
Model 1: Baseline characteristics	1.26**	1.59**	1.40**	1.00**	1.32**	2.20**	1.37**	1.08**	1.24**	1.57**	1.06**	0.75**
Model 2: Add time-varying variables			1.33**	0.87**							1.03**	0.70**
n (unique cases)	483	483	845	845	615	615	1,251	1,251	875	875	1,733	1,733

Notes: OLS = ordinary least squares regression models; Heckman = Heckman two-stage selection models; RE = random-effects time-series regression models; FE = fixed-effects time-series regression models. Complete case estimates eliminate cases using listwise deletion. Model 2 using complete cases with mother-reported variables is not estimated, since there is only one time-varying mother variable. All models include father, mother, and child baseline characteristics described in Table 3. Time-varying characteristics include fathers' work hours, father has a new partner, and mother has a new partner.

$^{\dagger}p < .10$; $^{*}p < .05$; $^{**}p < .01$

The random- and fixed-effects results yield large and statistically significant effects of coparenting on spending time, even when controlling for the time-varying variables. Again, the fixed-effects estimate is smaller than the random-effects estimate, indicating that part of the effect is attributable to characteristics that vary between cases; controlling for the time-varying characteristics yields point estimates of .85 and .70 for the random- and fixed-effects models (effect sizes of .74 and .61), respectively.

The pattern of results is very similar for the outcome of the frequency that the father engages in activities with the child: coparenting has a strong positive association with engagement. The Heckman estimate is slightly larger than the OLS estimate, the Mills' ratio is never statistically significant, the fixed-effects estimate is slightly smaller than the random-effects estimate, and adding time-varying covariates reduces the size of the effect (5%–13%). Effect sizes across all estimates range from .49 (for the fixed-effects result in Model 2) to .89 (for the Heckman model using the one-year data)—medium to large effects. Overall, the first panel of estimates suggests that there is strong evidence that cooperative coparenting is associated with all three measures of fathers' involvement—the number of days that nonresident fathers see their child at all, and the frequency with which fathers spend time and engage in activities with the child when they do have some contact.

To check for the robustness of the estimates described above, we repeated the analyses using our second sample. The middle panel of Table 4 shows the same series of results using the analytic sample with complete information on mother-reported variables. These analyses enable us to evaluate whether our results in the first panel were biased by using only cases in which the father was interviewed, since the fathers who participated in the survey are more highly connected to the mothers and more committed to the child than fathers not interviewed. These results present a very similar picture to those using complete case data for all variables: although the magnitudes of the coefficients sometimes change slightly (either smaller or larger), the overall association between coparenting and fathers' involvement is strong and statistically significant across all models for all three outcomes.

Finally, the last panel of Table 4 shows results using MI techniques to impute missing values for the full sample. Using MI enables us to include all nonresident father cases in which the mother was interviewed, the child lived with the mother, and the father saw the child at least once since the previous survey. As would be expected, we find a smaller association between coparenting and the number of days the father sees the child when we include the imputed data. There is likely more variability in the levels of coparenting and fathers' involvement (and a weaker link between the two) for imputed cases, since these pieces of information are imputed independently. Also, the imputed cases do not have the shared variance in the coparenting and fathers' involvement measures that arises from using the same individual (the mother) to report about both. Yet, the overall pattern in the results with imputed data is not challenged: greater cooperative coparenting is positively associated with fathers' seeing the child, spending time with the child, and engaging in activities with the child.

Moderating Factors

In order to evaluate whether there were differences in how coparenting affects fathers' involvement across subgroups, we reestimated our random-effects models and included interaction terms. We used the pooled sample with the random-effects model to maximize sample size. For each of the fathers' involvement outcomes, we interacted coparenting with child gender, the race/ethnicity variables, and father/mother having a previous child. There were no significant interactions for child gender. For the other two moderators, the only significant interactions were for the outcome of the number of days the father saw the child; these results are shown in Table 5. Coparenting has a marginally significant, smaller effect on the frequency the father sees the child for non-Hispanic white fathers, compared with non-Hispanic black fathers. In other words, contrary to expectations,

Table 5. Moderators of Coparenting on Nonresident Father Involvement

Number of Days Saw Child Past Month (including 0s)	β	*SE*
Father Race/Ethnicity (ref. = black non-Hispanic)		
Coparenting	5.96**	0.30
White non-Hispanic × Coparenting	–1.49†	0.79
Hispanic × Coparenting	–0.77	0.61
Other non-Hispanic × Coparenting	0.09	1.34
Child by Previous Partner		
Coparenting	6.02**	0.40
Father has another child × Coparenting	–1.13*	0.50
Mother has another child × Coparenting	0.49	0.50

Notes: Models are estimated using random effects and include the father, mother, and child baseline characteristics described in Table 3, as well as time-varying variables for father's work hours, father has a new partner, and mother has a new partner. Missing covariates are imputed using multiple imputation.

†$p < .10$; *$p < .05$; **$p < .01$

coparenting appears to be slightly *more* important for African American fathers' ties to children compared with white fathers.

Also, coparenting is less strongly linked to father-child contact if the father has a child by a previous partner than if he does not have any previous children by others; there is no difference by whether the mother has a previous child. Both of the significant interactions are relatively modest in magnitude, reducing the size of the main effect by 19%–25% but not eliminating the significant association between coparenting and father-child contact for the group being tested. For neither spending time nor engaging in activities is there any significant difference in how coparenting affects fathers' involvement by subgroup. Overall, we conclude that regardless of child gender, paternal race/ethnicity, and parents' previous fertility, more effective coparenting promotes fathers' involvement, even if the effects on father-child contact are slightly smaller for white fathers or when fathers have a previous child by another partner.

Control Variables

We include a number of control variables in order to try to reduce spuriousness in our estimates of how coparenting affects fathers' involvement. Since these variables are not our main focus, we show estimates in Appendix Table A1 for the covariates for two models predicting the number of days the father saw the child. The Heckman model simultaneously estimates the two equations predicting nonresidence and the number of days, conditional on nonresidence; we show both sets of coefficients. We also show results on the covariates from the random-effects model that pools nonresident fathers across waves.

With respect to the factors that affect becoming a nonresident father, we find that white and Hispanic fathers are significantly less likely to live away from their children than black fathers, and immigrant fathers are less likely to be nonresident than native-born fathers. Parents having a greater number of children together deters living apart, while fathers who have a previous child by another partner are much more likely to be nonresident at Year 1. Pro-fathering attitudes reduce the likelihood of becoming nonresident, while more-religious fathers are slightly more likely to be living apart. Having a substance problem, being physically violent, and having a history of incarceration each predict a greater chance of the father being nonresident at Year 1.

Turning to the predictors of the number of days per month that the nonresident father saw the child, we find that the direction of the effects of the covariates is generally similar across the Heckman model and the random-effects model, although there are some differences in which covariates reach statistical significance. We find that nonresident immigrant fathers see their children about two fewer days per month than fathers born in the United States; this may be because these fathers are spending more time out of the country. Fathers' age and education are not significantly related to the frequency of father-child contact. Fertility history is an important factor for fathers' seeing children: when the parents have other biological children together, the nonresident father is more likely to stay connected to the focal child in the pooled sample (but not the one-year sample). By contrast, when the father has a child by a previous partner, he sees the child an average of 2.5 to 4.1 fewer days per month. Fathers who have a history of incarceration see their child between two and three fewer days per month; we suspect that fathers who have spent time in jail/prison may be less capable or invested in performing the father role and that mothers may discourage such fathers from remaining involved with the child. Child characteristics do not appear to affect the frequency of father-child contact. With respect to the time-varying characteristics, consistent with the literature on "swapping" families (Manning and Smock 1999), when fathers go on to have a new partner, they are much less likely to see the focal child. Similarly, mothers' having a new partner has an even greater effect on the biological fathers' involvement with the child ($p < .01$). In other words, when the "package deal" that links fathers' partner and parental roles around a given child (Furstenberg and Cherlin 1991; Townsend 2002) comes apart, fathers appear to lose connection to their child(ren).[7]

DISCUSSION

In this paper, we examined the prevalence of fathers' involvement after a nonmarital urban birth and analyzed how cooperative coparenting between mothers and nonresident fathers is linked to fathers' involvement in children's lives. We focus on nonresident fathers because the majority of unwed couples who have children will be living apart within only a few years of their baby's birth. We find that 63% of fathers who bear a child outside of marriage will be living apart from the mother and child by the time the child is five years old, figures that are similar but slightly higher than those calculated from the 1979 National Longitudinal Survey of Youth (NLSY; Lerman and Sorenson 2000). The slightly higher fraction in the Fragile Families sample could be because the NLSY is a national sample rather than an urban sample (where father absence is more common). The difference could also be due to cohort differences because the births in the NLSY occurred during the 1980s, whereas the births in the Fragile Families Study occurred during the late 1990s. Taken together, these two studies suggest that about three-fifths of children born outside of marriage will experience parenting by a custodial mother and nonresident father by around age 5. Yet, little is known about these family dynamics and the extent to which parents' ability to cooperate may have long-term consequences for child well-being.

With respect to the level of involvement among nonresident fathers, we find notable divergence in the extent to which fathers remain involved over time. While some nonresident fathers remain significantly involved—seeing and spending time with the child and regularly engaging in father-child activities—a sizable fraction of fathers appear to have little connection to their children. By the time the child was age 5, 37% of nonresident fathers had not seen their child at any point over the previous two years, a figure in range

7. We do not show results for the covariates predicting the frequency of spending one or more hours or engaging in activities. In general, we find that when we limit the sample to those fathers who saw their child more than once in the previous month, few of the control variables are statistically significant; fathers' having a child by a previous partner is the most consistent (negative) predictor across models; none of the demographic characteristics is ever more than marginally significant.

with those in a recent study that compared four national studies showing that between one-fifth and three-fifths of nonresident fathers had not seen their child in the past year (Argys et al. 2007). By contrast, 43% of nonresident fathers saw their child more than once in the previous month, a figure somewhat lower than Lerman and Sorenson's (56% in the past month) about four years after the child was first observed in the survey (Lerman and Sorenson 2000). Again, these differences could be due to differences in the sample or the cohort of parents, as well as to the fact that our measure is seeing the child *more than* once per month, while theirs was seeing the child *at least* once per month.

The fact that a large fraction of five-year-old children born outside of marriage have no regular contact with their nonresident biological father is disquieting, since research has increasingly pointed to the benefits of high-quality involvement by nonresident fathers for children's well-being (Amato and Gilbreth 1999; Carlson 2006; King and Sobolewski 2006). At the same time, a large number of unwed fathers *do* have regular contact at child's age 5, perhaps more than would have been expected based on the divorce literature showing low levels of nonresident fathers' involvement (Cherlin 1992; Furstenberg, Morgan, and Allison 1987). Also, some evidence suggests that nonresident fathers become more involved in certain ways (such as having conversations) as children get older (Cooksey and Craig 1998). Yet, since the biggest drop-off in involvement for (mostly divorced) nonresident fathers occurs five years after the couple's romantic relationship ends (Seltzer 1991), we may expect greater paternal disengagement among the Fragile Families fathers in the future as well (given that the couple relationships have mostly been dissolved for less than five years at the time of the five-year survey). Considering these factors conjointly, we expect there may be greater variation in fathers' involvement over time, and the antecedent factors may differentially predict which fathers become more involved versus which become less involved. Fathers with greater social, psychological, and economic resources may become more involved at the same time that less advantaged fathers disengage, further reifying the inequality of resources that children receive over time (McLanahan 2004).

We find quite a high degree of coparenting among custodial mothers and nonresident fathers of young children during the five years after they have a nonmarital birth. Among all fathers who had seen the child at least once between surveys (the minimum threshold for mothers to be asked the coparenting questions), the average reported level of coparenting falls just above the midpoint of the 1-to-3 scale, and the level is higher for fathers who had recent contact with the child. These figures are notably higher than those reported by Sobolewski and King (2005) in their study of (mostly) divorced parents with adolescent children: they found, for example, that the mean score for how often the mother and father discussed the child was only 2.4 on a 1-to-6 scale. We suspect that coparenting among our never-married parents may be higher for several reasons. First, the Fragile Families mothers are asked the coparenting questions only if the father had at least minimal contact in the intervening period, so the certain zeros (i.e., no contact = no coparenting) are not included as they seem to be in the NLSY data. Second, the children in our study are significantly younger (age 5, compared with age 10–18), and less time has passed since the parents' relationship dissolved (the NLSY youth in the Sobolewski and King study had lived away from their fathers an average of 9.8 years). Third, relationship dissolution among unwed parents may be less of a traumatic "break" than a divorce, which requires legal action, may be highly conflictive, and typically ends a relationship of longer duration (and hence brings a greater sense of loss).

This research extends the literature on coparenting and nonresident fathers' involvement—which has previously focused primarily on divorced fathers—to focus explicitly on the role of fathers in children's lives following a nonmarital birth. Consistent with Sobolewski and King (2005), we find that the degree to which nonresident couples can cooperate in rearing their child encourages fathers to remain involved. Indeed, when the mother trusts the father and can communicate with him about the child's needs, the father is more likely to see the child at all, and to spend time and engage in activities with the child more frequently when he does have contact. The size of the effects ranges from small to large, depending on the particular estimation technique; smaller effects are observed in the structural equation models that allow for feedback effects between coparenting and fathers' involvement than in the regression models that posit a unidirectional effect. Our paper extends previous research by considering coparenting and nonresident fathers' involvement as dynamic constructs that change over time and that may be reciprocally related. We explicitly evaluate both whether coparenting affects fathers' involvement and whether fathers' involvement also influences coparenting. We find strong evidence for effects going from coparenting to fathers' involvement and only weak evidence for effects going in the opposite direction.

An obvious question is whether coparenting has a *causal* effect on nonresident fathers' involvement, or whether simply the type of men who are able to effectively coordinate parenting with mothers are also more likely to remain involved with their children *ex ante.* Our research is particularly instructive in this regard, since we use several different techniques designed to address the issue of unobserved heterogeneity. We find remarkable consistency in our substantive findings about the association between coparenting and fathers' involvement across various techniques, and across the one-year and pooled samples, regardless of whether we use complete cases or the full sample of nonresident fathers with imputed data on missing covariates.

Consistent with Sobolewski and King (2005), we find few significant differences by subgroup in how coparenting is linked with fathers' involvement. Only the father's having children by a previous partner is a strongly significant moderator, and the small magnitude of the interaction term does not negate the overall positive association: coparenting promotes father-child contact when fathers have a child by a previous partner, but it has a *stronger* effect when fathers do not have any children by other partners.

Several limitations of our research should be noted. First, we recognize that fathers may be involved in other ways that we do not measure here. For example, we do not measure any communication from fathers to children from afar, such as telephone calls, cards/letters, or e-mail. Further, our analyses do not examine fathers' economic contributions, which may complement or substitute for direct involvement. Recent research shows that about one quarter of nonresident fathers pay formal child support to their children three years after a nonmarital birth, and even more fathers make financial contributions outside the formal child support enforcement system (Nepomnyaschy 2007). How coparenting relates to economic support is a topic for additional investigation.

A second limitation concerns our use of mothers' reports of fathers' involvement with children. Using mothers' reports allows us to have information about all fathers, even those not interviewed. It is by now well-known that most nationally representative data sets underrepresent fathers, particularly those who live apart from their children (Garfinkel et al. 1998; Lerman 1993; Seltzer and Brandreth 1995). However, mothers may not have accurate information about the frequency and content of nonresident fathers' involvement with children (Coley and Morris 2002; Seltzer and Brandreth 1995), and the extent of their knowledge is likely correlated with the degree of cooperative coparenting. Also, using mothers' reports about both coparenting and fathers' involvement may inflate the observed correlations, since the same respondent could be over- or underreporting positive feelings of all kinds, sometimes referred to as "shared method variance" (Marsiglio et al. 2000). When we examined fathers' reports of father involvement for the subset of interviewed nonresident fathers (results not shown), we found that reports on the two fathers' involvement outcomes asked of both mothers and fathers (number of days and engagement in activities) are significantly and moderately correlated (at .21–.51 depending on measure and wave); also, reestimating the models using fathers' reports yielded results that are similar to the OLS, random-effects, and fixed-effects models using mothers' reports, although the magnitude of the coefficients is slightly smaller.[8] Hence, our results do not seem to be solely driven by the shared method variance inherent in using mothers' report for both coparenting and fathers' involvement.

A final limitation concerns inference of causality. Even with our extensive efforts to address selection bias in our estimates, we recognize that survey data are inherently inferior to experimental design for identifying a causal effect. Thus, we must be careful in interpreting our findings, as unobserved variables could be causing both cooperative coparenting and the fathering outcomes. Our fixed-effects models offer the most rigorous test of causality, since they control for time-invariant individual differences by focusing on within-subject change. Yet, these models do not account for unmeasured time-varying characteristics, so our results could still be biased by variables correlated with both coparenting and fathers' involvement that are changing over the observation period; also, these models do not account for the reciprocal nature of the association between coparenting and father involvement.

We conclude that many nonresident fathers are involved in their children's lives in the years following a nonmarital birth. This is encouraging, given the important role for fathers in the lives of their children. At the same time, the children in our study are young, and even by this early stage of children's development, many nonresident fathers are no longer in regular contact. Fathers who have lost touch completely may be unlikely to reengage later on, and involvement by those who remain connected at child's age 5 could diminish over time as more fathers go on to have additional children with new partners. This is disconcerting because an extensive literature suggests that, on average, father absence is disadvantageous for children (Amato 2005; Cherlin 1999; McLanahan and Sandefur 1994). While there is growing evidence for the benefits of high-quality involvement by nonresident fathers for child well-being (Amato and Gilbreth 1999; King and Sobolewski 2006), the relatively low prevalence of highly involved nonresident fathers suggests that such involvement will do little to obviate the consequences of father absence for children at the population level (Carlson 2006).

Our results also have implications for current policy efforts toward unmarried couples with children. The Healthy Marriage Initiative is now developing programs designed to strengthen couple relationships in order to facilitate getting (and staying) married for couples who so choose. To the extent that these efforts are successful in promoting stable marriages, fewer children may live away from their fathers in the first place. Yet, we suspect that even with a genuinely successful intervention that achieves healthy, stable marriages for some couples, many unwed couples will still break up soon after their baby's birth. To the extent that cooperative coparenting is enhanced indirectly with relationship skills training focused on the couple's relationship and communication, such programs may facilitate fathers' remaining connected to their nonresident children once the couple relationship has ended. At the same time, policy interventions might usefully focus not only on strengthening the couple's romantic relationship but also explicitly on strengthening their ability to work together in rearing their child, particularly when couples have (or will have) children by other partners; early evidence from one intervention among low-income couples suggests that such programs hold promise for strengthening fathers' ties to children (Cowan et al. 2007). Since coparenting has been shown to be an important element of family life that is distinct from both couple relationship quality and parenting, incorporating coparenting into such curricula could enhance childrearing for all parents, whether living in the same or separate households.

8. The Heckman results are never statistically significant.

Appendix Table A1. Coefficients From Heckman Selection Models and Random-Effects Models

	Heckman Model[a]				Random-Effects Model[b]	
	Nonresidence (0/1)		Number of Days		Number of Days	
Variables	β	*SE*	β	*SE*	β	*SE*
Baseline Characteristics						
Age at baby's birth (mean)	–0.01	0.00	0.06	0.05	0.04	0.03
Race/ethnicity (ref. = black non-Hispanic)						
White non-Hispanic	–0.49**	0.09	–2.55†	1.33	–0.11	0.65
Hispanic	–0.38**	0.07	–0.88	0.96	–0.21	0.48
Other non-Hispanic	–0.18	0.16	–1.48	1.68	–0.69	1.03
Immigrant (born outside the U.S.)	–0.24*	0.12	–2.32†	1.26	–1.80*	0.70
Father lived with both parents age 15	–0.02	0.08	1.24	0.82	0.65	0.42
Education (ref. = less than high school)						
High school	0.12†	0.06	–0.64	0.72	0.46	0.42
Some college or more	0.08	0.08	–0.88	0.87	0.16	0.50
Number of biological children with mother	–0.14**	0.04	0.44	0.46	0.75**	0.23
Child(ren) by other partner	0.29**	0.06	–4.11**	0.78	–2.45**	0.40
Mother has child(ren) by other partner	–0.07	0.06	–0.20	0.65	0.27	0.41
Pro-fathering attitudes (mean; range = 1–4)	–0.30**	0.07	0.82	0.93	0.49	0.43
Religious attendance (mean; range = 1–5)	0.06*	0.02	–0.27	0.30	–0.16	0.16
Self-reported health (mean; range = 1–5)	0.03	0.03	–0.26	0.36	–0.26	0.20
Substance problem	0.35**	0.13	–1.98	1.20	–1.21†	0.72
Physical violence toward mother	0.36**	0.13	–0.53	1.19	–0.91	0.72
History of incarceration	0.15*	0.06	–2.83**	0.68	–2.10**	0.38
Baby is a boy	0.08	0.05	0.13	0.60	0.52	0.36
EASI "difficult" temperament (mean; range = 1–5)	–0.02	0.03	–0.11	0.27	–0.10	0.17
Mills' ratio (lambda)			0.51	2.71	NA	
Time-Varying Characteristics						
Work hours per week (mean)	NA		NA		0.01†	0.01
Father has a new partner	NA		NA		–1.54**	0.41
Mother has a new partner	NA		NA		–2.76**	0.31
Number of Unique Cases (*n*)	2,650		1,228		2,191	

Note: Missing covariates are imputed using multiple imputation.

[a]The Heckman selection model estimates the two equations simultaneously; the full sample at Year 1 (2,650) is used to predict nonresidence, and the nonresident cases (1,228) are used to predict number of days with the inverse probability of selection (Mills' ratio) included as an additional variable. This model corresponds to column j, top row in Table 4.

[b]The random-effects model uses the pooled sample of nonresident fathers across survey Years 1, 3, and 5. This model corresponds to column k, top row in Table 4.

†$p < .10$; *$p < .05$; **$p < .01$

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