

Longitudinal Relations Between Coparenting and Father Engagement in Low-Income Residential and Nonresidential Father Families

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Coparenting relationship quality and father involvement are closely linked but few studies have investigated this relationship using samples of socioeconomically disadvantaged families. The current study used family systems theory to examine the longitudinal and bidirectional relations between coparenting relationship quality and father engagement in caregiving and play, using a large and racially diverse sample of low-income residential and nonresidential fathers in the Building Strong Families project ($N = 1,908$). Structural equation modeling tested cross-lagged relations between couple-level coparenting and father engagement at two time points for both residential and nonresidential father families. For residential fathers, positive coparenting at 15 months predicted father engagement in caregiving at 36 months. There was no support for a bidirectional or unidirectional model between coparenting and father engagement in play for either residential or nonresidential fathers. There were significant concurrent relations between coparenting and father engagement in caregiving and play for both residential and nonresidential fathers, providing support for positive spillover in line with family systems theory.

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Approximately 40% of children are now born to unmarried couples in the United States (Martin, Hamilton, Osterman, Driscoll, & Matthews, 2017). This group is becoming a sizable population in the country, yet we know little about family processes that shape parenting among unmarried couples. In recent years, researchers and policymakers alike have expressed growing interest in low-income unmarried parents, and significant resources have been invested to ensure family stability and healthy child outcomes for these families (Brown, 2010; Wood, Moore, Clarkwest, & Killewald, 2014). As part of the United States

Department of Health and Human Services' Healthy Marriage and Responsible Fatherhood initiative, Congress approved \$150 million per year for 5 years in 2005. The funding continued in 2010 with an additional \$75 million per year to fund marriage education programs that would increase the likelihood that low-income couples marry and fathers remain residential and involved in the lives of their children (Wood et al., 2014). These efforts explicitly acknowledged that coparenting and father involvement were important family processes that would benefit both couples and their children. As such, the current study focused on examining the relations between positive coparenting relationship quality and father engagement in caregiving and play in a large sample of socioeconomically disadvantaged couples with young children from the Building Strong Families (BSF) program evaluation, using two longitudinal time points within the first 3 years after the birth of a new child.

Defining Father Involvement and Father Engagement in Caregiving and Play

Father involvement has been operationalized in a number of different ways (Lamb, Pleck, Charnov, & Levine, 1985). According to Lamb et al. (1985)'s tripartite model of father involvement, there are three components: (a) availability (i.e., fathers are present and available for interaction with the child), (b) responsibility (i.e., fathers are responsible for making arrangements for children's daily activities, such as childcare and doctor's visits), and (c) direct engagement (i.e., fathers are directly involved in positive interactions with the child). In the current study, we focused on fathers' direct engagement in caregiving and cognitively and socially stim-

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ulating play activities. The focus on father engagement in caregiving and play is supported by prior research with low-income fathers that has demonstrated the effects of fathers' engagement in these activities for promoting children's development.

For instance, results from the Early Head Start Research and Evaluation Project demonstrated that low-income fathers' supportive parenting during toy play was positively linked with toddlers' receptive vocabulary and emotion regulation (Cabrera, Karberg, Malin, & Aldoney, 2017; Cabrera, Shannon, & Tamis-LeMonda, 2007). When low-income fathers used explanatory language and creative play in positive interactions with their 24-month-olds, children were five times more likely to fall within the normative cognitive range on the Bayley Scales of Infant Development compared with children with less positive father-child interactions (Shannon, Tamis-LeMonda, London, & Cabrera, 2002). Finally, children whose fathers took on 40% or more of the caregiving tasks had better academic outcomes compared to children whose fathers were less involved (Halle, 2002).

Family Systems Theory and Relations Between Coparenting and Father Engagement

Family systems theory (Cox, Paley, & Harter, 2001) posits that families are comprised of interrelated subsystems—including the mother-father, mother-child, and father-child relationships—that exert direct and indirect influence on each other. The mother-father subsystem often includes interactions unrelated to parenting, such as romantic relationship quality, as well as interactions that are related to parenting, such as coparenting. Coparenting refers to the manner in which parents work together in their roles as parents (Feinberg, 2003). It is an ongoing, interactive, cooperative, and mutually supportive relationship that is primarily focused on the child and child-rearing activities in contrast to the romantic relationship between partners. Coparenting has been examined in both residential and nonresidential father families, although how one defines coparenting and father engagement for families in which the father is residential versus nonresidential is an important consideration that few studies address (Fagan, Kaufman, & Dyer, 2019; Fagan & Palkovitz, 2011).

Interdependence among family subsystems is a key tenet of family systems theory (Cox et al., 2001), which suggests that the mother-father and father-child subsystems have reciprocal influences on each other (Fagan & Cabrera, 2012). Poor coparenting relationship quality can “spill over” and negatively affect the father-child relationship, which, in turn, can adversely affect children's developmental outcomes (Erel & Burman, 1995; Krishnakumar & Buehler, 2000). Arguments between coparents can undermine positive father-child interactions and play a role in whether or not fathers remain connected and involved in the family. This is particularly salient for nonresidential fathers because access to their child may be partly dependent on mothers, who typically have child custody and may engage in maternal gatekeeping (Carlson, McLanahan, & Brooks-Gunn, 2008). If mothers perceive fathers as unsupportive or hostile coparents, parents may not work cooperatively as a team to coparent the child, and in the end, mothers may discourage father involvement (i.e., maternal gatekeeping) or fathers may simply stay away to avoid further conflict. Positive coparenting relationship quality can also spill over and lead to more supportive father-child relation-

ships (Perry, Harmon, & Leeper, 2012). Mothers' positive views and support of the father as a coparent can encourage fathers' involvement and their competence in the parenting role, which may enhance positive father-child relationships and, in turn, promote healthy developmental outcomes for young children (Fagan & Palkovitz, 2011).

Interdependence in the family system also means there can be reciprocal effects such that the quality of the father-child relationship may predict the quality of the mother-father relationship. When fathers expend more effort and time engaging in parenting activities, such as caregiving and play with their children, the coparenting relationship between mothers and fathers may flourish (Carlson et al., 2008). Conversely, uninvolved or hostile father engagement may contribute to mothers' negative perceptions of fathers as coparents, gradually leading to further distance and conflict between mothers and fathers. Overall, scholars have noted that the mother-father relationship and parenting practices cannot be fully understood without addressing bidirectional, reciprocal relations (Krishnakumar & Buehler, 2000), which we do in this study.

Because of the social and political interests in responsible fatherhood and promotion of marriage education, as well as the fact that children in nonresidential father families are at risk for poorer social, emotional, cognitive, and academic outcomes (Brown, 2010), we focused on both the concurrent and longitudinal relations between coparenting and the father engagement in residential and nonresidential father families. Understanding these linkages could reveal the connections between these family subsystems and whether supportive coparenting relations at one point predicted later father engagement with children or vice versa. Gaining a better understanding of these directional relations between coparenting and fathering can help pinpoint targets for future intervention efforts.

Prior Evidence Supporting Relations Between Coparenting and Father Engagement

Prior research examining bidirectional, longitudinal relations between coparenting and father engagement in low-income families has found mixed results, with some finding that coparenting when children were 1 to 3 years old predicted nonresidential fathers' engagement with children when they were 3 to 5 years old (Carlson et al., 2008; Fagan & Palkovitz, 2011, 2019), using data from the Fragile Families and Child Wellbeing Study. Yet, Fagan and Cabrera (2012), using data from the Early Childhood Longitudinal Survey-Birth Cohort, found that residential fathers' engagement (i.e., caregiving and cognitive stimulation) with 9-month-old infants was a stronger predictor of coparenting conflict at 24 and 48 months than vice versa.

One limitation of earlier work is that few studies have investigated longitudinal, bidirectional relations between coparenting and father engagement in low-income families with young children, and even fewer studies have examined such models using samples of residential and nonresidential fathers (for exceptions, see Fagan & Palkovitz, 2011, 2019). Such investigations are needed if we want to understand how processes within different family ecologies either inhibit or promote low-income fathers' engagement with their children (Volling et al., 2019). The mechanisms linking coparenting and father engagement may be different for low-

income residential- and nonresidential-father families simply by virtue of whether the fathers are living with the mother and child or not. Residential fathers have greater physical access to their children, more opportunities for direct engagement, and are better able to coordinate coparenting responsibilities with the mother on a daily basis compared to nonresidential fathers, who may have tenuous ties with the mother and child due to limited physical contact and engagement. In the end, residential and nonresidential fathers are likely to spend different amounts of time and engage in different activities with their children in addition to their ability to participate in everyday decision making related to coparenting (Carlson et al., 2008).

Another limitation of prior work is how coparenting has been conceptualized and analyzed. Often, statistical models are conducted separately for mothers and fathers rather than considering the dyadic nature of coparenting and utilizing maternal and paternal reports to create a latent construct of coparenting. By utilizing a latent dyad approach (Gonzalez & Griffin, 2001) in our structural equation modeling (SEM), we created a latent variable that reflected the shared variance of coparenting across maternal and paternal reports, and then used this shared component in our SEM models to predict latent variables of both fathers' caregiving and play.

The Current Study

The primary objective of the current study was to examine concurrent and bidirectional longitudinal relations between coparenting and father engagement among low-income residential and nonresidential father families with young children. Family processes may very well differ between these families based on what it means to be an engaged residential or nonresidential father. Using the family systems theory, we hypothesized positive relations between coparenting quality and father's engagement in caregiving and play, as well as bidirectional relations between coparenting and father engagement for both residential and nonresidential fathers, but also considered that there would be stronger relations between coparenting and father engagement for residential than nonresidential fathers by virtue of living together with the mother and child and having more opportunities than nonresidential fathers to coordinate coparenting responsibilities and interact with their child.

Method

The BSF Project

Mothers and fathers were participants in the BSF project, a large-scale randomized controlled trial of healthy marriage and relationship education interventions for low-income unmarried couples (Wood, McConnell, Moore, & Clarkwest, 2010). Because BSF included longitudinal follow-ups at 15 and 36 months after couples enrolled in the intervention, we were able to examine longitudinal relations between coparenting and father engagement over time in a cross-lagged panel model. One of the strengths of BSF is that information about coparenting and fathering was collected directly from fathers and mothers rather than relying strictly on mothers' reports of father engagement (Carlson et al., 2008). Supplemental material S1 (available online) shows the

conceptual model tested in the current study, demonstrating the bidirectional, cross-lagged paths between coparenting and father engagement (caregiving and play) across 15 and 36 months.

Participants

Participants for the current study were 1,908 BSF families. Families were enrolled in the BSF intervention when mothers were pregnant or shortly after the BSF child's birth. During the follow-up interviews at 15 and 36 months, fathers were asked whether they lived with the mother and child, 1 (*none of the time*), 2 (*some of the time*), 3 (*most of the time*), or 4 (*all of the time*). Consistent with several prior studies examining residential fathers (Fagan, Levine, Kaufman, & Hammar, 2016; Waller & Dwyer Emory, 2014), we defined consistently residential fathers as those fathers who reported living with the mother and child *all or most* of the time at both Time 2 (T2) and Time 3 (T3; $n = 1,499$ fathers). We also created a subgroup of consistently nonresidential fathers by focusing on men who reported living with the mother and child *some or none* of the time at both T2 and T3 ($n = 409$). Sociodemographic information of both groups is displayed in Table 1.

Procedure

The BSF project developed, implemented, and evaluated healthy marriage and relationship education programs designed to strengthen the relationships of low-income, unmarried couples who were expecting or recently had a baby (Wood et al., 2010). The project was funded by the Office of Planning, Research and Evaluation in the Administration for Children and Families, United States Department of Health and Human Services and conducted by Mathematica Policy Research from 2005 to 2011. The project recruited 5,102 opposite-sex couples from hospitals, maternity wards, prenatal clinics, health clinics, and special nutritional programs for Women, Infants, and Children (WIC) from eight sites in the U.S. Couples were eligible to enroll if (a) both the mother and father agreed to participate in the intervention, (b) they were romantically involved, (c) they were either expecting a baby together or had a baby younger than 3 months old, (d) they were unmarried at the time the baby was conceived, and (e) both parents were 18 years and older (Wood et al., 2010). Mathematica Policy Research obtained participants' written consents and randomly assigned couples into an intervention group ($n = 2,553$) or a control group ($n = 2,549$).

Given the aim to strengthen low-income unmarried couples' relationships, the BSF intervention focused primarily on providing 30 to 42 hr of relationship skills education in the form of group sessions, with each session ranging from 2 to 5 hr depending on the day of the week. The intervention group received adapted versions of existing relationship education curricula (e.g., "Love's Cradle"; "Loving Couples, Loving Children"; "Becoming Parents") and were taught relationship skills, including developing commitment and trust, managing relational conflict, and navigating multiple-partner fertility (defined as a parent having biological children with more than one partner; Monte, 2017; Wood et al., 2014). Couples also received individual support from family coordinators, who reinforced the intervention's relationship skills curriculum, provided emotional support, encouraged couples' participation in

Table 1
 Descriptive Statistics: Sociodemographic and Key Variables by Father Residential Status

Variable and coding	Full sample (<i>N</i> = 1,908), <i>M</i> (<i>SD</i>) or <i>n</i> (%)	Consistently residential (<i>n</i> = 1,499), <i>M</i> (<i>SD</i>) or <i>n</i> (%)	Consistently nonresidential (<i>n</i> = 409), <i>M</i> (<i>SD</i>) or <i>n</i> (%)	<i>p</i> value
Age in years	25.79 (6.32)	26.12 (6.35)	24.60 (6.12)	<.001
Number of biological children with mother	1.37 (.85)	1.41 (.87)	1.25 (.77)	<.001
Race and ethnicity				<.001
1 = Non-Hispanic White	309 (16.52%)	267 (18.16%)	42 (10.50%)	
2 = Non-Hispanic Black	1,014 (54.22%)	702 (47.76%)	312 (78%)	
3 = Non-Hispanic other	33 (1.76%)	27 (1.84%)	5 (1.50%)	
4 = Hispanic	514 (27.49%)	474 (32.24%)	40 (10%)	
Education				<i>ns</i>
0 = Does not have high school diploma	603 (31.89%)	478 (32.19%)	125 (30.79%)	
1 = Has high school diploma or equivalent	1,177 (62.24%)	920 (61.95%)	257 (63.30%)	
2 = Other	111 (5.87%)	87 (5.86%)	24 (5.91%)	
Working for pay (yes)	1,473 (77.28%)	1,193 (79.69%)	280 (68.46%)	<.001
Total earnings in the past year				<.001
0 = None	101 (5.79%)	73 (5.36%)	28 (7.35%)	
1 = \$1-\$4,999	295 (16.92%)	195 (14.32%)	100 (26.25%)	
2 = \$5,000-\$9,999	250 (14.34%)	192 (14.10%)	58 (15.22%)	
3 = \$10,000-\$14,999	340 (19.51%)	274 (20.12%)	66 (17.32%)	
4 = \$15,000-\$19,999	298 (17.10%)	254 (18.65%)	44 (11.55%)	
5 = \$20,000-\$24,999	218 (12.51%)	191 (14.02%)	27 (7.09%)	
6 = \$25,000-\$34,999	150 (8.61%)	116 (8.52%)	34 (8.92)	
7 = \$35,000 or above	91 (5.22%)	67 (4.92%)	24 (6.30%)	
Received welfare in the past year (yes)	905 (47.43%)	733 (48.90%)	172 (42.05%)	<.05
Father has child(ren) with another partner (yes)	569 (29.92%)	431 (28.79%)	138 (34.07%)	<.05
Couple's relationship length in years	3.50 (3.35)	3.69 (3.46)	2.81 (2.77)	<.001
BSF project group status (intervention)	972 (50.94%)	761 (50.77%)	211 (51.59%)	<i>ns</i>
Mother's report of coparenting at Time 2	4.34 (.83)	4.58 (.50)	3.40 (1.17)	<.001
Father's report of coparenting at Time 2	4.52 (.61)	4.67 (.40)	3.96 (.89)	<.001
Mother's report of coparenting at Time 3	4.25 (.94)	4.54 (.54)	3.04 (1.23)	<.001
Father's report of coparenting at Time 3	4.50 (.65)	4.67 (.41)	3.86 (.93)	<.001
Father caregiving at Time 2	5.211 (1.01)	5.40 (.81)	4.36 (1.34)	<.001
Father play at Time 2	4.51 (1.08)	4.68 (.98)	3.76 (1.18)	<.001
Father caregiving at Time 3	5.01 (1.06)	5.21 (.90)	4.02 (1.24)	<.001
Father play at Time 3	4.50 (.98)	4.66 (.87)	3.71 (1.10)	<.001

Note. Unless otherwise indicated, all sociodemographic variables are from baseline. Significant tests are between residential and nonresidential groups. Chi-square tests are used for categorical variables and ANOVAs are used for continuous variables. BSF = Building Strong Families; *ns* = nonsignificant.

group sessions, and offered needs assessments and referrals to additional services (e.g., employment, mental health). The control group did not receive these services but could seek relationship skills education from sources other than BSF. Overall, the goal of the BSF project was to strengthen low-income couples' relationships so that they could create and sustain a stable and healthy home environment for their children (Wood et al., 2014).

In the BSF project, there were three time points in which data collection occurred: baseline (Time 1 [T1]), the 15-month follow-up (T2), and the 36-month follow-up (T3). At T1, mothers and fathers completed an eligibility survey and enrolled in the project. Two follow-up telephone interviews were conducted with both parents at approximately 15 and 36 months after enrollment. According to BSF documentation, children were approximately 15 months old at T2 and 37 months old at T3 (Wood et al., 2010). Only data from T2 and T3 were analyzed in the current study because these were the time points at which all families had children, and coparenting and father engagement measures were collected. Most women were still pregnant (61.54%) at T1. Because the BSF project team found no effects of the intervention on key outcomes, including couples' relationship quality, coparenting, likelihood of marriage, and father involvement (Wood et al.,

2014), intervention and control families were combined for the current analyses, and BSF randomization status was included as a control variable. The Institutional Review Board (IRB)—Health Sciences and Behavioral Sciences at the (University of Michigan)—determined that secondary analysis of BSF data was exempt from IRB oversight.

Measures

Coparenting relationship quality and father engagement were measured at T2 and T3. Two latent variables of fathers' engagement were created using items reflecting caregiving and cognitive and social play.

Coparenting relationship quality. During individual telephone interviews at T2 and T3, mothers and fathers were asked to report on the quality of the coparenting relationship, using 10 items from the Parenting Alliance Inventory (PAI; Abidin & Brunner, 1995). The PAI is a well-established self-report measure that taps into parents' cognitive and emotional evaluations of the coparenting relationship (Palkovitz, Fagan, & Hull, 2013). Items focused on parental communications related to the child (e.g., "My child's other parent and I communicate well about our child") and coparenting

renting solidarity or the supportive alliance between coparents (e.g., "I feel good about my child's other parent's judgment about what is right for our child"). The measure required that mothers report on the father as a coparent, and fathers report on the mother as a coparent. Both mothers and fathers rated the 10 items on a 5-point scale (1 = *strongly agree* to 5 = *strongly disagree*), which was reversed so that higher scores reflected more positive coparenting. Cronbach's alphas were 0.95 for mothers and 0.94 for fathers at T2, and 0.97 for mothers and 0.93 for fathers at T3.

Fathers' engagement in caregiving. During telephone interviews at T2 and T3, fathers were asked to report on the frequency of their engagement in caregiving activities using three items (i.e., "Helped child to get dressed," "Changed child's diapers or helped him or her use the toilet," "Given child a bottle or something to eat"). These items have been used frequently in other large-scale studies (e.g., National Evaluation of Early Head Start, Fragile Families and Child Wellbeing Study; Wood et al., 2010). Fathers were instructed to rate the frequency with which they engaged in these activities in the past month on a 6-point scale (1 = *more than once a day* to 6 = *not at all*), which was reversed so that higher scores reflected more engagement in caregiving activities. Reliabilities were $\alpha = 0.86$ and $\alpha = 0.74$ at T2 and T3, respectively.

Father's engagement in cognitive and social play. Fathers were also asked during interviews at T2 and T3 to report on the frequency of their engagement in cognitively and socially stimulating play activities within the past month using four items (i.e., "Sung songs with child," "Read or looked at books with child," "Told stories to child," "Played with games or toys with child"). These items have also been frequently used in other large-scale studies (e.g., National Evaluation of Early Head Start). Fathers used the same 6-point scale (1 = *more than once a day* to 6 = *not at all*), which was reversed so that higher scores reflected more engagement in cognitive and social play. Reliabilities were $\alpha = 0.78$ and $\alpha = 0.81$ at T2 and T3, respectively.

Sociodemographic control variables. Ordinary least squares regressions were conducted to search for potential covariates. Results revealed five sociodemographic variables related either to coparenting or father engagement: (a) fathers' ethnicity/race, (b) fathers' education, (c) fathers' work status, (d) fathers' multiple-partner fertility, and (e) and participation in the BSF intervention. Non-Hispanic Black fathers reported significantly lower levels of assessments of their child's mother as a coparent at T2, $\beta = -0.08$, $p < .05$, and higher levels of caregiving at T2, $\beta = 0.09$, $p < .05$, compared to non-Hispanic White fathers. Hispanic fathers reported significantly lower assessments of mothers as coparents at T2, $\beta = -0.16$, $p < .001$, and T3, $\beta = -0.09$, $p < .01$, less caregiving at T2, $\beta = -0.12$, $p < .01$, and T3, $\beta = -0.09$, $p < .01$, and less play at T3, $\beta = -0.11$, $p < .01$, compared to non-Hispanic White fathers. Fathers with a high school diploma reported significantly better coparenting relations at T2, $\beta = 0.18$, $p < .01$, compared to fathers with no high school diploma, and employed fathers reported significantly less caregiving, $\beta = -0.05$, $p < .05$, and play at T3, $\beta = -0.07$, $p < .01$, compared to unemployed fathers. Fathers with higher numbers of children with more than one partner had significantly lower assessments of mothers as coparents at T3, $\beta = -0.07$, $p < .05$; less caregiving at T2, $\beta = -0.08$, $p < .01$, and T3, $\beta = -0.08$, $p < .01$; and less play at T2, $\beta = -0.09$, $p < .001$, compared to those with lower numbers of children with more than one partner.

The BSF intervention group reported significantly more positive coparenting relations at T2, $\beta = 0.07$, $p < .01$, and T3, $\beta = 0.07$, $p < .01$, and significantly less caregiving at T2, $\beta = -0.09$, $p < .01$, compared to the control group. Couples' relationship length, a potential proxy for romantic relationship quality (Shafer, Jensen, & Larson, 2014), was also added as a control variable based on prior research showing relations with coparenting and father involvement (Jia & Schoppe-Sullivan, 2011). Couples' relationship length was also significantly associated with father caregiving such that longer relationship length was significantly related to less father caregiving at T2, $\beta = -0.09$, $p < .01$. In sum, we included fathers' ethnicity/race, education, work status, multiple-partner fertility, couples' relationship length, and participation in the BSF intervention as covariates in our main analyses.

Model Development and Data Analysis Plan

Preliminary analyses and data reduction. Given the nature of the longitudinal data involving both maternal and paternal reports, preliminary analyses involved confirmatory factor analysis (CFA) to build latent constructs of coparenting and father engagement at T2 and T3 separately before proceeding to model latent variables at both times. Additionally, we tested for measurement invariance using the BSF randomization group status¹ and fathers' residential status as separate grouping variables to determine whether latent constructs were measurement invariant across groups (Kline, 2016). If the same factor structure fits the data across groups equally well, then the measurement model was deemed to have configural invariance (Little, 2013) and other stricter levels of measurement invariance could be tested subsequently.

In general, testing individual CFA models separately aligns with a model-building approach where individual models are built incrementally, with each subsequent model built on the previous model. Testing a multivariate model without such a model building approach can lead to misspecification in one part of the model that is masked by good fit elsewhere, rendering the overall model to falsely fit the data. Building the model of interest from the smallest specified pieces ensured that all the pieces in the model were appropriately specified and fit the data well (Kline, 2016).

Missing data. Stata's (Version 14; StataCorp, 2015) missingness pattern analysis and logistic regressions were used to examine missing data. Stata's missingness pattern analysis showed that data were missing in 0% to 3.07% (for mothers' reports of coparenting

¹ To ensure that we could combine the BSF control and intervention groups, we conducted measurement invariance testing using the control and intervention status as the grouping variable. We first examined configural invariance versus metric invariance. The Satorra-Bentler scaled chi-square difference test between the configural and metric invariance models was significant for the coparenting and father caregiving model, $\Delta\chi^2(44) = 64.50$, $p < 0.05$, and coparenting and father play model, $\Delta\chi^2(46) = 62.84$, $p < 0.05$, indicating that configural invariance was present across control and intervention groups but not metric invariance. More advanced levels of measurement invariance could not be tested without achieving metric invariance. Although we could not fully achieve measurement invariance between the BSF control and intervention groups, because previous work found no effects of the BSF intervention status on main variables of interest, including coparenting relationship quality and father engagement (Wood et al., 2014), we decided to combine the two groups. This allowed us to create sufficient sample sizes for our analyses. We also included BSF intervention status as a control variable to err on the side of caution.

at T3) of the cases. Data for mothers' reports of coparenting at T2 were missing in 2.40% of the cases and 0.27% and 0.20% of the cases for fathers' reports of coparenting at T2 and T3, respectively. Data for father engagement in caregiving at T2 and T3 were missing in 0.47% and 0.20% of the cases, respectively. Finally, data for father engagement in play at T2 and T3 were missing in 0.27% and 0.47% of the cases, respectively. Across all control variables, data were missing in less than 0.13% of cases with the exception of couples' relationship length, which had missing data in 1.40% of the cases.

Results from logistic regressions demonstrated that missing cases for father play at T2 were missing at random (MAR), where missing values were significantly associated with fathers' reports of coparenting at T2 ($p = .031$). Missingness of all other key predictor and outcome variables (e.g., mothers' reports of coparenting, fathers' reports of coparenting, father engagement in caregiving) was not related to any of the observed variables in the dataset, suggesting they were missing completely at random (MCAR). The actual missing data mechanism was more likely to be MAR given the possibility that missing cases in these key predictor and outcomes variables depended on observed variables in the original BSF dataset and not the truncated dataset for the current analyses. To account for all cases and missing data patterns, we used full information maximum likelihood (FIML), which estimates parameters by maximizing the sample and using all available data (Kline, 2016). FIML has been shown to produce less biased and more efficient estimates than other missing data methods (i.e., listwise deletion, mean imputation), especially when data do not appear to be MCAR (Allison, 2003).

Building latent variables. The latent variables were built first around the larger sample of residential fathers and then subsequently tested with nonresidential fathers by conducting measurement invariance testing using fathers' residential status as the grouping variable. Initial CFAs focused on using individual items to create first-order latent variables of coparenting and father engagement that could be used in the subsequent SEM models. Because each parent reported on the other parent's coparenting rather than their own coparenting, both mothers' and fathers' first-order coparenting latent variables were used to create a second-order couple-level latent variable to assess the dyadic nature of the coparenting construct (Gonzalez & Griffin, 2001). Individual CFAs to assess model fit for the latent variables of mothers' and fathers' reports of coparenting separately, as well as the second-order couple-level coparenting latent variable at T2 and T3 were conducted (for model fit indices, see [online supplemental material S2](#), available online).

In the second-order couple-level coparenting latent variable models, we fixed the loadings for mothers' and fathers' reports of coparenting to be equal to each other at 1. We also fixed the residual variances of mothers' and fathers' reports of coparenting to be equal. These constraints ensured that mothers' and fathers' reports were contributing equally to the dyadic latent variable. Individual CFA models had good fit at T2 and T3. Second-order couple-level coparenting CFA models also had good fit at T2 and T3. These results suggested that the mothers' and fathers' reports of coparenting cohered adequately to form the anticipated second-order couple-level coparenting latent variable. The CFA model combining the second-order couple-level coparenting latent variable at T2 and T3 also demonstrated good fit.

For the father engagement measures, we used fathers' reports of engagement in various caregiving and play activities in the CFA models to create the latent variables of father caregiving and father play and examined the fit for these latent variables at T2 and T3. Father caregiving CFA models at T2 and T3 were just identified because there were only three indicators per model. Father play CFA models had good fit at T2 and T3. We ran CFA models combining the second-order couple-level coparenting and father's engagement latent variables at T2. The model had good fit. We ran identical models for T3 and found good fit as well. Finally, we assessed a measurement model combining second-order couple-level coparenting and father's engagement latent variables at T2 and T3. The model had good fit. Factor loadings for all latent variables can be found in [online supplemental material S4](#).

Building cross-lagged models. [Online supplemental material S1](#) (available online) shows the conceptual bidirectional models guiding the SEM analyses. The first model estimated both the stability paths for couple-level coparenting and father caregiving across T2 and T3, as well as bidirectional paths between coparenting and father caregiving across the two time points. An identical second model was conducted when father play was the indicator of father engagement. In all the models, cross-sectional covariances between latent variables were included. SEM was conducted using the R package lavaan (Version 0.6–2; Rosseel, 2012) to estimate the models. The robust maximum likelihood estimation method was used because the data did not meet the multivariate normality assumption based on Mardia's test.

Model fit was evaluated using several fit indices (Kline, 2016), including root-mean-square error of approximation (RMSEA; Steiger, 1990; <0.06 for good fit), 90% confidence interval (CI) of RMSEA (Kenny, 2015; <0.05 for lower bound for good fit), comparative fit index (CFI; Bentler, 1990; >0.95 for good fit), and standardized root-mean-square residuals (SRMRs; Hu & Bentler, 1999; <0.05 for good fit). To compare cross-lagged paths in some of the models, we used a Satorra-Bentler chi-square test (Satorra & Bentler, 2001), comparing an unconstrained model in which all paths were allowed to be freely estimated to a constrained model in which the cross-lagged paths were constrained to be equal. A statistically significant chi-square test indicates that the cross-lagged paths are significantly different from each other and that one should prefer the unconstrained model over the constrained model (Kline, 2016). We also relied on the Akaike information criterion (AIC; Akaike, 1974) for further confirmation of our results. The model with the smaller AIC is preferred over that of the larger AIC.

Results

Preliminary Analyses

We first tested for measurement invariance of the coparenting and father engagement latent variables across residential and nonresidential father families. The residential status of the father was used as the grouping variable. We first ran a configural invariance model and then a metric invariance model. The two models were compared using the Satorra-Bentler scaled chi-square difference test. Tests results were significant for the coparenting and father caregiving CFA model, $\Delta\chi^2(44) = 710.56, p < .001$, and the coparenting and father play CFA model, $\Delta\chi^2(42) = 205.75, p <$

.001. These results indicated that configural invariance was present, meaning that we were able to fit the same set of indicators to the coparenting and father engagement latent variables for both residential and nonresidential fathers, but not metric invariance. The fact that metric invariance did not hold means that we were not able to impose equality constraints on factor loadings for the two groups.

Configural, metric, and scalar invariance should be sequentially established to obtain measurement invariance (Kline, 2016). Without achieving measurement invariance, conducting multigroup analysis to compare structural models and studying differences between groups is not fully meaningful (Milfont & Fischer, 2015), as establishing measurement invariance is a prerequisite for meaningful comparisons across groups. Given that stricter measurement invariance tests (i.e., scalar invariance) cannot be conducted without achieving metric invariance first, our coparenting and father engagement latent variables were deemed noninvariant across groups, and as a result, we were unable to conduct a formal multigroup analysis of structural models that would allow us to empirically compare paths of interests between residential and nonresidential fathers groups. Instead, we ran separate cross-lagged panel models for residential and nonresidential fathers. Sociodemographic variables at T1 that showed significant relations with the main study variables were used as covariates in all the models.

Low-Income Residential Father Models

Coparenting and father caregiving model. For the first model examining coparenting and father caregiving (Figure 1), the estimation converged normally, and the model had good fit to the data, $\chi^2(1359) = 2,719.26$, $p < .001$, RMSEA = 0.03, 90% CI

[0.03, 0.03], CFI = 0.95, SRMR = 0.03. The stability paths for coparenting and father caregiving at T2 and T3 were both significant: coparenting, $B = 0.65$, $SE = 0.13$, $\beta = 0.63$, $p < .001$, 95% CI [0.46, 0.80], father caregiving, $B = 0.55$, $SE = 0.06$, $\beta = 0.48$, $p < .001$, 95% CI [0.38, 0.57]. Results from the Satorra-Bentler scaled chi-square test comparing the unconstrained model to the constrained model in which the coparenting to father caregiving path and father caregiving to coparenting path were constrained to be equal was statistically significant, $\chi^2(1) = 5.16$, $p = .023$, indicating that the models were significantly different and the cross-lagged paths differed from each other. The AIC for the unconstrained model was also smaller (AIC = 93,419) than that of the constrained model (AIC = 93,423), suggesting that the unconstrained model should be preferred over the constrained model. The path from coparenting at T2 predicting father caregiving at T3 was statistically significant, $B = 0.77$, $SE = 0.37$, $\beta = 0.14$, $p = .037$, 95% CI [0.02, 0.25], whereas the path from father caregiving at T2 to coparenting at T3 was not statistically significant, $B = 0.01$, $SE = 0.01$, $\beta = 0.06$, $p = .245$, 95% CI [-0.04, 0.17]. Cross-sectional covariances between coparenting and father caregiving were only significant at T2. Overall, results indicated that positive coparenting at T2 predicted an increase in father caregiving at T3, controlling for earlier levels of father caregiving, and that this path significantly differed from the alternate bidirectional path between father caregiving at T2 to coparenting at T3.

Coparenting and father play model. For the second model examining coparenting and father play (see the online supplemental material S3 Figure 2A), the estimation converged normally, and the model had good fit to the data, $\chi^2(1468) = 2,926.75$, $p < .001$, RMSEA = 0.03, 90% CI [0.03, 0.03], CFI = 0.95, SRMR = 0.03. There were significant stability coefficients for coparenting from

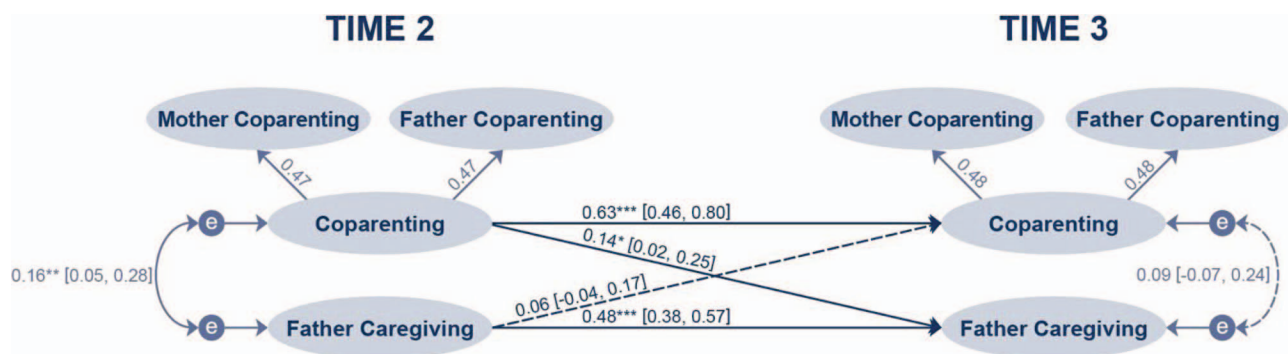


Figure 1. Results from the first cross-lagged model of couple-level coparenting relationship quality and fathers' engagement in caregiving amongst low-income residential fathers with standardized parameter estimates. $\chi^2(1359) = 2719.26$, $p < .001$, RMSEA = 0.03, 90% CI [0.03, 0.03], CFI = 0.95, SRMR = 0.03. The model controlled for father's ethnicity/race, education, work status, multiple-partner fertility, couple's relationship length, and BSF intervention status. Being Hispanic ($\beta = -0.29$, $p < .001$), having a high school diploma ($\beta = 0.11$, $p < .05$), having other education ($\beta = 0.10$, $p < .05$), and being in the BSF intervention group ($\beta = 0.11$, $p < .05$) were significantly associated with couple-level coparenting at T2. Being non-Hispanic Black ($\beta = 0.10$, $p < .01$), being Hispanic ($\beta = -0.14$, $p < .001$), multiple-partner fertility ($\beta = -0.09$, $p < .01$), and couple's relationship length ($\beta = -0.10$, $p < .01$) were significantly associated with father's engagement in caregiving at T2. Being non-Hispanic Black ($\beta = 0.12$, $p < .05$) was significantly associated with couple-level coparenting at T3. No socio-demographic control variables were significantly associated with father's engagement in caregiving at T3. Dotted lines indicate nonsignificant paths. * $p < .05$. ** $p < .01$. *** $p < .001$. See the online article for the color version of this figure.

T2 to T3, $B = 0.66$, $SE = 0.13$, $\beta = 0.63$, $p < .001$, 95% CI [0.45, 0.81], and for father play from T2 to T3, $B = 0.61$, $SE = 0.05$, $\beta = 0.61$, $p < .001$, 95% CI [0.54, 0.67]. Neither of the cross-lagged paths between coparenting and father play was statistically significant, with coparenting at T2 to father play at T3, $B = 0.15$, $SE = 0.28$, $\beta = 0.03$, $p = .578$, 95% CI [-0.08, 0.15], and father play at T2 to coparenting at T3, $B = 0.01$, $SE = 0.01$, $\beta = 0.02$, $p = .689$, 95% CI [-0.09, 0.13]. Cross-sectional covariances between coparenting and father play were significant at both T2 and T3.

Low-Income Nonresidential Father Models

Coparenting and father caregiving model. For the first model examining coparenting and father caregiving (Figure 2), the estimation converged normally, and the model had moderate fit to the data, $\chi^2(1360) = 2,663.93$, $p < .001$, RMSEA = 0.05, 90% CI [0.05, 0.05], CFI = 0.90, SRMR = 0.11. The stability paths for coparenting and father caregiving at T2 and T3 were both significant: coparenting, $B = 0.76$, $SE = 0.11$, $\beta = 0.76$, $p < .001$, 95% CI [0.58, 0.95], and father caregiving, $B = 0.41$, $SE = 0.08$, $\beta = 0.46$, $p < .001$, 95% CI [0.30, 0.62]. Neither the path from coparenting at T2 to father caregiving at T3, $B = 0.18$, $SE = 0.21$, $\beta = 0.09$, $p = .41$, 95% CI [-0.12, 0.30], nor the path from father caregiving at T2 to coparenting at T3, $B = -0.02$, $SE = 0.04$, $\beta = -0.05$, $p = .53$, 95% CI [-0.22, 0.11], was significant. Cross-sectional covariances between coparenting and father caregiving were significant at both T2 and T3.

Coparenting and father play model. For the second model examining coparenting and father play (see the online supplemental material S3 Figure 2B), the estimation converged normally, and the model had moderate fit to the data, $\chi^2(1468) = 2,858.08$, $p < .001$, RMSEA = 0.05, 90% CI [0.05, 0.05], CFI = 0.90, SRMR = 0.10. The stability paths for coparenting and father play at T2 and T3 were both significant: coparenting, $B = 0.66$, $SE = 0.12$, $\beta =$

0.67 , $p < .001$, 95% CI [0.47, 0.87], and father play, $B = 0.50$, $SE = 0.08$, $\beta = 0.55$, $p < .001$, 95% CI [0.39, 0.70]. Neither the path from coparenting at T2 to father play at T3, $B = 0.01$, $SE = 0.21$, $\beta = 0.00$, $p = .98$, 95% CI [-0.22, 0.23], nor the path from father play at T2 to coparenting at T3, $B = 0.07$, $SE = 0.05$, $\beta = 0.13$, $p = .21$, 95% CI [-0.07, 0.33], was significant. Cross-sectional covariances between the coparenting and father play were significant at both T2 and T3.

Discussion

According to family systems theory, the mother-father and father-child subsystems are interrelated, and there are mutually reciprocal, bidirectional relations between these subsystems over time. The main goal of this study was to test whether there was support for these bidirectional relations between coparenting and father engagement in a sample of low-income residential and nonresidential father families. We found evidence for both similarities and differences across the two groups. Regarding similarities, findings revealed that there was significant stability in coparenting and fathers' caregiving over time and, in general, significant within-time covariances between coparenting and fathers' caregiving for both residential and nonresidential fathers. Similarly, there was significant stability in coparenting and father play over time and significant within-time covariances between coparenting and father play for both groups of fathers. Regarding differences, the final measurement invariance results revealed that the latent variables of coparenting and father engagement—both caregiving and play—were noninvariant or different across residential and nonresidential father groups, indicating that these may very well be different constructs across these two family ecologies. This may help explain differences in regressions estimates found between residential and nonresidential fathers.

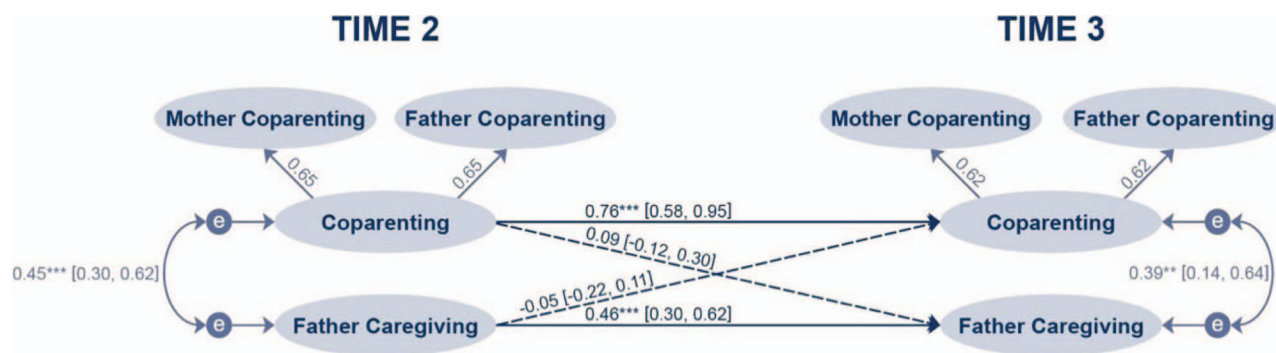


Figure 2. Results from the first cross-lagged model of couple-level coparenting relationship quality and fathers' engagement in caregiving amongst low-income nonresidential fathers with standardized parameter estimates. $\chi^2(1360) = 2663.93$, $p < .001$, RMSEA = 0.05, 90% CI [0.05, 0.05], CFI = 0.90, SRMR = 0.11. The model controlled for father's ethnicity/race, education, work status, multiple-partner fertility, couple's relationship length, and BSF intervention status. Being non-Hispanic Black ($\beta = 0.25$, $p < .01$), having education other than some high school education or a high school diploma ($\beta = -0.44$, $p < .05$), and couples' relationship length ($\beta = 0.03$, $p < .01$) were significantly associated with couple-level coparenting at T2. Being Hispanic ($\beta = -0.59$, $p < .05$) was significantly associated with father's engagement in caregiving at T3. No socio-demographic control variables were significantly associated with father's engagement in caregiving at T2 and couple-level coparenting at T3. Dotted lines indicate nonsignificant paths. ** $p < .01$. *** $p < .001$. See the online article for the color version of this figure.

The primary analyses tested for bidirectional relations between coparenting and father engagement in both caregiving and play. We found no support for bidirectional relations in either residential or nonresidential father groups. Only the longitudinal path for coparenting at 15 months significantly predicted father caregiving at 36 months for residential fathers. When mothers and residential fathers supported each other with open communication earlier in their relationship 15 months after the BSF intervention, fathers were engaged in more caregiving activities at 36 months. Positive coparenting quality in the current study appeared to be important for residential fathers' future engagement in caregiving. This makes sense, given that these men were consistently residential throughout the 21 months between the two time points, whereas nonresidential fathers were not. Family systems theory would predict that there is positive spillover from one family subsystem (e.g., mother-father) to another (e.g., father-child), and this appeared to be the case both within and across time for residential fathers living in the same space with their partner and children.

For nonresidential fathers, we only found significant within-time covariances between coparenting and father engagement variables at each time, suggesting that, similar to residential fathers, it was the quality of the concurrent coparenting relationship that was related to nonresidential fathers' engagement in both caregiving and play. Because nonresidential fathers may not be present consistently, what occurs at one point in time in the family may not predict what occurs at a subsequent time point. This lack of longitudinal prediction may be due to a number of interpersonal processes, including greater coparenting conflict and maternal gatekeeping in nonresidential father families. Mothers who perceive fathers as more argumentative with respect to parenting their children and less cooperative in their role as a coparent may limit fathers' access to and future involvement with their children (Carlson et al., 2008).

It should be noted that we were only able to make these observations about similarities and differences across residential father groups descriptively and not from direct statistical comparisons of the paths in our models because we did not find measurement invariance in our coparenting and father engagement CFA models between the two groups. Because measurement invariance is a necessary prerequisite for conducting multigroup analysis of structural models (Kline, 2016), we were unable to engage in direct empirical testing of cross-lagged paths across residential father groups. Relatedly, the lack of longitudinal prediction for the nonresidential father sample may have been due to the smaller sample size compared to that of residential fathers and the fact that they were not a homogeneous group (we had to combine fathers reporting they lived with their children some or none of the time for sufficient sample size). One must also consider how different the family circumstances are when fathers reside with the mother and child and when they do not. Together, these factors may have contributed to the results we found across residential and nonresidential fathers.

Overall, our results are consistent with Fagan and Palkovitz (2011, 2019), who found longitudinal links between coparenting when the child was 1 year old and father engagement when the child was 3 and 5 years old. However, our results are inconsistent, especially with Fagan and Palkovitz (2011), in that they found longitudinal links for nonresidential nonromantic father families but not for residential father families. The inconsistencies may be

attributed to a number of factors that differed across the studies. Fagan and Palkovitz (2011) created three distinct groups (i.e., residential, nonresidential romantic, and nonresidential nonromantic parents), whereas we only focused on residential and nonresidential fathers. Further, they analyzed mothers' and fathers' reports of coparenting separately, whereas we used mothers' and fathers' reports to create a dyadic latent variable. Finally, the researchers used path analysis with observed variables only, whereas our analyses involved SEM with latent variables.

Coparenting and Father Engagement in Caregiving and Play

Even though positive coparenting was related to father caregiving and play for both groups of fathers within time, we only found longitudinal prediction of residential fathers' caregiving. Over the years, fathers have increased the number of hours they devote to caring for their young children. The amount of time fathers invest in housework and childcare tasks has tripled between 1965 and 2016, with fathers spending, on average, 8 hr a week caring for their children in 2016 compared to 2.5 hr in 1965 (Parker & Livingston, 2019). There is also evidence to show that ethnic and racial minority fathers, especially non-Hispanic Black fathers, are highly involved in caregiving. Both non-Hispanic Black residential and nonresidential fathers have reported higher scores on feeding, bathing, diapering, and dressing their children under 5 years old in the past month compared to non-Hispanic White residential and nonresidential fathers (Jones & Mosher, 2013). Greater time devoted to caregiving for low-income minority men over the years may be one reason for the findings linking coparenting and father caregiving.

Specific to coparenting and father play, there was no evidence of bidirectional or longitudinal prediction for either residential father group. There were only significant within-time covariances. Both mothers and fathers engage in caregiving and play activities with their children. However, fathers spend relatively more time in play than caregiving compared to mothers and often prefer play, particularly physically stimulating rough-and-tumble play (Paquette, Carbonneau, Dubeau, Bigras, & Tremblay, 2003; Parke, 2013), over caregiving activities. Given their preference for and comfort with play activities, fathers may engage in this sort of enjoyable and playful interaction with their children regardless of coparenting relationship quality, which may explain why positive coparenting at one time did not necessarily predict fathers' engagement in play at a later time for both residential and nonresidential fathers.

Limitations and Considerations for Future Research

There are several limitations to note. Even though information on coparenting was obtained directly from both parents and father engagement from fathers themselves, BSF did not include extensive observational measures of coparenting or father-child interaction, which are often used in related research in early childhood (Feinberg, 2003; Jia & Schoppe-Sullivan, 2011). Future studies may benefit from a multimethod study design, where both parents' reports and observational measures of coparenting and father engagement are the sources of information. Another limitation involves the measure of coparenting, which did not assess negative

aspects of coparenting, such as conflict and undermining, which have been used repeatedly in other coparenting studies (Feinberg, 2003; Jia & Schoppe-Sullivan, 2011). Different findings may have emerged had negative dimensions of coparenting been available. This could have allowed us to address directly whether coparenting conflict is one reason why nonresidential fathers did not live with their partner and children in the first place and why we found few longitudinal, bidirectional relations between coparenting and father engagement in this group. Relatedly, because BSF did not include a measure of maternal gatekeeping, we were unable to address directly the role of this interpersonal process in understanding relations between coparenting and father engagement. We recommend that future research consider how these sorts of interpersonal relations between coparents encourage or discourage fathers' engagement with their children.

Our findings cannot be generalized to all low-income unmarried parents with young children because BSF data were collected from parents who volunteered to participate in an intervention to strengthen their couple relationship. Finally, limitations of the cross-lagged panel model must be acknowledged, as the autoregressive and cross-lag paths may reflect both within-person (trait-like, intraindividual characteristics that endure over time) and between-person (rank-order stability in individual differences) effects, and the cross-lagged modeling approach used here does not allow us to disaggregate within-person from between-person effects. Because longitudinal data are multilevel with time nested within individuals, including a random intercept into the cross-lagged panel model is one means of separating the within-person, traitlike stability from the between-person, temporal stability. However, such a model requires at least three waves of data and ideally lagged intervals that are equally spaced over time (Hamaker, Kuiper, & Grasman, 2015), which was simply not the case here. Future studies with more than two waves of data, which can move beyond the cross-lagged panel model analysis used here, are clearly needed to address this issue further.

Notwithstanding these limitations, a key strength of the current study was the focus on coparenting and father engagement in low-income families, and an examination of the bidirectional relations between the two for residential and nonresidential fathers. Our findings suggest that how one defines coparenting and father engagement for residential and nonresidential low-income fathers may differ, even though there were strong ties between coparenting and father's engagement in caregiving and play concurrently, even if not longitudinally, for both residential and nonresidential fathers. Thus, intervention efforts targeting low-income fathers may want to consider building the coparental alliance, particularly around daily responsibilities and childcare tasks, in order to encourage and maintain low-income father's involvement with their children (Pruett, Cowan, Cowan, Gillette, & Pruett, 2019).

Future research may also benefit from developing more nuanced ways of measuring fathers' residential status and undertake more ecologically valid models that incorporate change in fathers' residential status and family dynamics over time (Mitchell et al., 2015; Volling et al., 2019). Couple relationships dissolve, fathers move to another residence, new romantic relationships are formed, and additional children are born. As such, fathers may be residential in one household and nonresidential in another, but these sorts of changing social dynamics with multiple partners and with children of multiple partners are rarely examined in research on

fathering and family relationship functioning. Related future research with both residential and nonresidential fathers will be needed in order to address how these complex family dynamics contribute to coparenting relationship quality and the role of fathers in children's development.

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