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Family Structure Transitions and Child Development: Instability, Selection, and Population Heterogeneity

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Abstract

A growing literature documents the importance of family instability for child wellbeing. In this article, we use longitudinal data from the Fragile Families and Child Wellbeing Study to examine the impacts of family instability on children's cognitive and socioemotional development in early and middle childhood. We extend existing research in several ways: (1) by distinguishing between the number and types of family structure changes; (2) by accounting for time-varying as well as time-constant confounding; and (3) by assessing racial/ethnic and gender differences in family instability effects. Our results indicate that family instability has a causal effect on children's development, but the effect depends on the type of change, the outcome assessed, and the population examined. Generally speaking, transitions out of a two-parent family are more negative for children's development than transitions into a two-parent family. The effect of family instability is stronger for children's socioemotional development than for their cognitive achievement. For socioemotional development, transitions out of a two-parent family are more negative for white children, whereas transitions into a two-parent family are more negative for Hispanic children. These findings suggest that future research should pay more attention to the type of family structure transition and to population heterogeneity.

Keywords

family structure transitions; child development; instability; selection; population heterogeneity

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A growing body of research finds that family structure instability is negatively associated with children's wellbeing (for a recent review, see Waldfogel, Craigie, and Brooks-Gunn 2010). Notably, Wu and Martinson (1993) found that family instability during childhood increased the risk of nonmarital childbearing, whereas living in a stable, single-mother family had no such effect. Since then numerous other researchers have documented an association between family structure change and declines in children's wellbeing, including declines in cognitive development (Magnuson and Berger 2009), increases in behavioral

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problems (Cavanagh and Huston 2006; Fomby and Cherlin 2007; Osborne and McLanahan 2007), and declines in physical health (Bzostek and Beck 2011).

A critical question is whether all changes in family structure are equally important or whether certain types of changes are more important than others. Whereas earlier studies suggest that all changes are equally harmful (Cavanagh and Huston 2006; Fomby and Cherlin 2007; Osborne and McLanahan 2007; Wu 1996), more recent work suggests that changes involving the exit of a parent from the household are more harmful for maternal wellbeing and children's behavioral outcomes, whereas changes involving the entrance of a parent or parent figure are more detrimental for children's achievement-related outcomes (Magnuson and Berger 2009; Mitchell et al. 2015; Osborne, Berger, and Magnuson 2012). Understanding how the effects of family instability vary by type of change and by outcome examined is important not only for disentangling how children's cognitive and socioemotional development is shaped by their family environment, but also for developing policies for improving child wellbeing. For example, recent programs aimed at promoting marriage among low-income, unmarried parents (e.g., Wood et al. 2012) may increase children's exposure to the entrance of a father or father figure into the household and, in some instances, lead to subsequent transitions if the relationship is unsuccessful. To anticipate the potential costs and benefits of these programs, we need a better understanding of how children respond to different types of family structure transitions.

A second question is whether the association between family structure instability and child wellbeing is causal. Unstable families likely differ from stable families in many ways, both measured and unmeasured, and researchers have used a variety of approaches to adjust for factors that select people into different family types, including models that control for lagged dependent variables, individual and family fixed-effects models, natural experiments, and instrumental variables methods (for a recent review, see McLanahan, Tach, and Schneider 2013). However, when researchers are dealing with multiple treatments occurring at multiple time points (e.g., the exit of a biological father followed by the entrance of a social father), standard panel models are inadequate to provide unbiased estimates of treatment effects. For example, the exit of a father at time 1 may lead to a decline in family income at time 2, which in turn may lead to the entrance of a new father figure at time 3. In cases such as this, alternative panel models are needed to address the dynamic relationship between family structure instability and time-varying covariates such as income.

A final question involves differences in children's responses to family instability across population subgroups. Generally speaking, previous research suggests that family instability effects are weaker (less negative) for racial/ethnic minority groups than they are for whites, and stronger (more negative) for boys than they are for girls (Cooper et al. 2011; Wu and Thomson 2001). To account for these differences, scholars have argued that minorities may respond less negatively to family instability than do whites because they have developed better support systems for dealing with such changes (McLoyd et al. 2000). Similarly, researchers have argued that because family transitions largely involve the gain or loss of a male role model, boys are more likely than girls to be negatively affected by such transitions (Allison and Furstenberg 1989). There is also some evidence that subgroup differences in children's responses to family structure instability may depend on the outcome being studied, with white children showing greater sensitivity when the outcome is behavior problems (Fomby and Cherlin 2007; Osborne and McLanahan 2007), and black children showing greater sensitivity when the outcome is cognitive/academic achievement (Heard 2007).

In this article, we draw on data from the Fragile Families and Child Wellbeing Study (FFCWS) to examine the impacts of family structure instability on children's cognitive and socioemotional development. Our study seeks to extend existing research in three ways. First, we examine family instability effects using both the number and the type of family structure transitions, which allows us to determine if instability effects differ by children's initial status. Second, we examine whether the effects of family instability on children's wellbeing are sensitive to different models that account for the processes that sort people into different family types and mediate the effects of family change. We compare estimates from random-effects models, fixed-effects models, and marginal structural models. As far as we know, our study is the first to use marginal structural models to examine the effects of family structure transitions on child outcomes. Random-effects models adjust for measured covariates; fixed-effects models adjust for unmeasured time-constant covariates; and marginal structural models adjust for measured time-varying covariates with multiple treatments. Finally, we consider population heterogeneity in the effects of family structure transitions on child outcomes by examining differences by race/ethnicity and by child gender. Together, our analysis provides a rigorous assessment of how different types of family structure transitions affect the cognitive and socioemotional development of children during early to middle childhood.

THE ROLE OF FAMILY INSTABILITY

Numerous studies find that children who experience a change in family structure lag behind children who grow up in stable family structures across multiple outcomes in different domains (Fomby and Cherlin 2007; Wu 1996). The instability hypothesis draws on the family stress model and posits that disparities in child outcomes arise because of the stress induced by reconfigurations in family composition. According to this view, family disruptions are often accompanied by changes in the roles and routines of parents and children alike, which in turn are associated with fluctuating parental resources, deteriorating parenting quality, and emotional insecurity (Coleman, Ganong, and Fine 2000; George 1993). In short, family structure instability likely undermines processes of effective socialization by parents.

Consistent with this perspective, a burgeoning body of research documents the adverse effects of family instability on a diverse set of child-related outcomes, including maternal wellbeing, parenting quality, cognitive and socioemotional development, and premarital birth (Beck et al. 2010; Brown 2006; Carlson and Corcoran 2001; Cavanagh and Huston 2006; Fomby and Cherlin 2007; Osborne and McLanahan 2007; Ram and Hou 2003; Wu 1996). Child development during early and middle childhood is of particular significance. Research on the life course and human capital formation identifies early to middle childhood as a critical and sensitive period in child development, meaning that children's developmental trajectories are the most malleable during this period and, once shaped, may be difficult to reverse at later life stages (Elder 1998; Shonkoff and Phillips 2000). Scholars have also shown that the cognitive and socioemotional skills developed during childhood are strong predictors of life course outcomes, such as academic achievement, health, educational attainment, labor market performance, and union formation (Heckman 2007). Thus, it is important to examine the links between family instability and early to middle childhood development.

Most prior research focuses on the *quantity* of change, that is, the number of family structure transitions a child experiences. For example, Wu and Thomson (2001) suggest that the frequency and intensity of family change are more consequential than the type of family change (e.g., divorce versus remarriage) for child and adolescent outcomes such as early sexual initiation. Similarly, Fomby and Cherlin (2007:183) contend that “the nature of the transition in terms of changes in household composition is less relevant than the stress associated with moving from one form to another.”

More recent research, however, suggests that the *type* of transition may be more important than the number of transitions (Magnuson and Berger 2009; Meadows, McLanahan, and Brooks-Gunn 2008; Osborne et al. 2012; Ryan and Claessens 2013). These studies indicate that the consequences of family instability likely differ depending on whether the change involves a move into or out of a two-parent family. The transition from single motherhood into a coresidential union may benefit children insofar as it increases their access to parental resources (time and money) and kinship support. At the same time, such a move may be harmful to children insofar as the entrance of a parent or parent figure disrupts family routines and may lead to conflict in parent-child relationships. Similarly, while the transition out of a two-parent family into a single-mother family is expected to be detrimental for children by reducing parental resources, children may benefit from such a move if the parental relationship involved high levels of conflict and if the mother has access to social support from her extended family.

Although the relative effects of different types of family structure transitions and the direction of effects are theoretically ambiguous, several studies find that different types of family structure transitions have differential impacts on child wellbeing. There is also evidence that the relative importance of entrances versus exits depends on the outcome being examined. On the one hand, exit of a biological parent from the household has negative effects on maternal wellbeing, such as material hardship and mental health, and children’s behavioral development (Magnuson and Berger 2009; Meadows et al. 2008; Mitchell et al. 2015; Osborne et al. 2012). On the other hand, entrance of a biological parent or parent figure appears to affect maternal wellbeing positively but children’s cognitive development negatively (Magnuson and Berger 2009; Osborne et al. 2012).

In summary, our review of existing research underscores the need to examine both the number and the type of family structure transitions in assessing family instability effects. This integrative approach is necessary to disentangle the extent to which the stress induced by family structure transitions depends on origin-destination patterns.

SELECTION INTO FAMILY INSTABILITY

Research on the effects of family instability on child wellbeing must confront the issue of selection bias, which occurs when the factors leading to family structure change are also associated with the child outcome of interest. To date, scholars have focused primarily on selection due to unobserved variables that do not change over time, applying child and family fixed-effects models (Aughinbaugh, Pierret, and Rothstein 2005; Dunifon and Kowaleski-Jones 2002; Foster and Kalil 2007; Hao and Xie 2002). These models identify the causal effect of family instability by comparing the difference in child outcomes before and after the event, or by comparing differences in the outcomes of siblings, one of whom experienced a transition and the other of whom did not. Studies using these approaches generally find that family structure transitions have a causal effect on child wellbeing, although the effects become smaller in magnitude than those estimated from conventional regression approaches (McLanahan et al. 2013). These findings are most robust for outcomes that measure children's socioemotional adjustment.

Even in instances where unmeasured covariates are not seen as a problem, researchers who want to estimate the causal effect of family instability on child wellbeing may face a dilemma with respect to how to treat measured covariates that change over time and are related to both family instability and child wellbeing. The life course perspective suggests that sequences of events and responses to events are reciprocally intertwined with one another (Elder 1985, 1998; Elder, Johnson, and Crosnoe 2003). The effects of time-varying changes in family structure, therefore, cannot be viewed in isolation but must be considered in the context of other time-varying factors. Consider maternal employment, which is associated with child outcomes as well as family structure. A mother's prior family structure change may affect her current employment status insofar as transitioning out of a coresidential union may increase her likelihood of working. A mother's current employment status, in turn, may affect subsequent family structure transitions, insofar as working may affect her likelihood of finding a new partner by increasing her exposure to men with more favorable labor market characteristics. This example describes how family structure transitions can affect as well as be affected by time-varying covariates, such as changes in employment. Family income is another time-varying covariate associated with child wellbeing that is both a cause and a consequence of family structure instability. The key point here is that time-varying covariates may confound and at the same time mediate the impact of time-varying family structure transitions on children's development.

This insight from the life course perspective on the time-varying nature of the correlates of family instability presents major challenges for modeling its effects, even in cases where covariates are measured (Elwert and Winship 2010; Sampson, Sharkey, and Raudenbush 2008). On one hand, if the researcher excludes time-varying covariates (e.g., maternal employment or family income), the model will overestimate the causal effect of family instability on child wellbeing. On the other hand, if the researcher includes time-varying covariates, the model will bias the causal effect of family instability through its inability to distinguish confounding from mediation.[1] In short, existing approaches to estimating the effect of multiple changes in family structure (i.e., multiple treatments) are unlikely to sufficiently handle the bidirectional relationship between time-varying family structure transitions and time-varying covariates.

Our study adopts a rigorous approach to addressing bias due to the presence of both unmeasured time-constant covariates and measured time-varying covariates. We begin by estimating a random-effects model (REM) that examines the association between family instability and child outcomes. To account for potential confounding by unmeasured time-constant characteristics, we apply a child fixed-effects model (FEM) that estimates family instability effects using within-child variation in family transitions and child outcomes across time (Dunifon and Kowaleski-Jones 2002; Foster and Kalil 2007). To account for potential bias due to measured time-varying characteristics, we use a marginal structural model (MSM). The model uses an inverse probability of treatment (IPT) weighting estimator by which children who experience family instability and those who do not are *sequentially* balanced on measured covariates (Robins 1999; Robins, Brumback, and Hernán 2000). As we will explain, the MSM differs from the REM in that it accounts for measured time-varying covariates that are both a cause and a consequence of family instability. Neither the FEM nor the MSM accounts for confounding by both unmeasured time-constant and measured time-varying covariates in the same model. However, using multiple panel models provides a more complete picture of the range of estimates than has been provided in past research.

HETEROGENEITY IN FAMILY INSTABILITY EFFECTS

The literature on family structure instability and child wellbeing has long acknowledged population heterogeneity. With respect to racial/ethnic differences, researchers have found that family instability effects are more salient for whites than for other racial/ethnic groups (Fomby and Cherlin 2007; Wu and Thomson 2001). One explanation for this difference is the degree of socioeconomic stress associated with family structure change (Fomby, Mollborn, and Sennott 2010). Compared to whites, racial/ethnic minority groups have encountered structural disadvantage in various forms and degrees, including persistent poverty, racial segregation, and low-quality schooling (Amato and Keith 1991; McLoyd et al. 2000; Wilson 1987). Any costs arising from moving out of a two-parent family may be attenuated for racial/ethnic minority children insofar as they are already dealing with multiple stressors engendered by higher levels of socioeconomic disadvantages. Similarly, any benefits arising from moving into a two-parent family may be dampened if such a move is not sufficient to move a family out of poverty. In this vein, the unique effects of family structure transitions may be smaller for racial/ethnic minority children than for white children.

A related explanation emphasizes the role of kinship networks in accounting for racial/ethnic differences in children's responses to family structure instability. Family instability and

[1]This model also creates collider stratification bias (Pearl 2009). Consider a scenario where unobserved factors affect time-varying covariates and child outcomes but not family instability. Even under the assumption of no unobserved confounding, conditioning on time-varying covariates invokes an unnecessary correlation between their common causes, that is, family instability and unobserved factors. Because unobserved factors also affect child outcomes, conditioning on time-varying covariates makes it impossible to distinguish the effect of family instability from that of unobserved factors.

complexity have increased across all racial/ethnic groups over the past half century, but these characteristics are more common among racial/ethnic minority groups than among whites. As a result, racial/ethnic minority groups have likely developed larger kinship networks for coping with family instability (Cherlin and Furstenberg 1986; McLoyd et al. 2000; Stack 1974). If children in racial/ethnic minority families are more likely than white children to rely on extended kinship networks, they may be better equipped to absorb the impacts of changes in the nuclear family structure. Sarkisian and Gerstel (2004) suggest that whereas financial and emotional support from kin are more available to white families, practical support is more available to racial/ethnic minority groups. Because practical support involves the physical presence of extended family members through help with childcare, housework, and transportation, it may be more effective for lessening the turbulence arising from family structure change. As a result, family instability effects may be dampened for racial/ethnic minority children.

With respect to gender differences, prior studies suggest that family instability has more deleterious consequences for boys than for girls (Cavanagh, Crissey, and Raley 2008; Cooper et al. 2011). Family structure transitions are typically characterized as the presence or absence of (or a change in) a male role model, which may be more consequential for boys' identity formation than for girls (Allison and Furstenberg 1989). Therefore, mothers' relationships with sons during family disruptions may undergo greater changes than their relationships with daughters (Hetherington, Cox, and Cox 1985). Moreover, because levels of cognitive and socioemotional development during early to middle childhood tend to be relatively lower for boys than for girls (Entwisle, Alexander, and Olson 1997), boys may adjust to family structure transitions at a slower pace than girls.

While invaluable, research to date has provided only limited insights on population heterogeneity in family instability effects. First, due to data limitations, most studies examine black-white differences; with few exceptions, little research explores the effects of family instability on Hispanics (Fomby et al. 2010). This omission is unfortunate because Hispanics represent the largest and fastest growing minority population in the United States. They also exhibit distinctive patterns of family processes; for example, coresidential unions are more common and more stable among Hispanics than among blacks (Osborne, Manning, and Smock 2007). Second, studies have yet to incorporate the number and the type of family transitions in assessing population heterogeneity in the effects of family instability on child development. Finally, with a few exceptions (e.g., Cooper et al. 2011), most of our knowledge about population heterogeneity is based on associational studies, which likely obscure family instability effects and selection effects.

DATA AND METHODS

Data

Data come from the Fragile Families and Child Wellbeing Study (FFCWS), a longitudinal birth cohort study of 4,898 children born between 1998 and 2000 in 20 U.S. cities with populations greater than 200,000 (Reichman et al. 2001). The FFCWS design called for an oversample of births to unmarried parents, yielding a sample in which a quarter of the births were to married parents and three-quarters were to unmarried parents. Baseline interviews were conducted shortly after the birth, with mothers interviewed in the hospital and fathers interviewed either in the hospital or by phone as soon as they could be located. Response rates for the baseline survey were 82 percent for married mothers, 87 percent for unmarried mothers, 89 percent for married fathers, and 75 percent for unmarried fathers.[2] Follow-up surveys were conducted when the focal child was 1, 3, 5, and 9 years of age. Response rates for the Years 1, 3, 5, and 9 surveys were 91, 88, 87, and 76 percent, respectively, for mothers who completed the baseline interview.

The FFCWS provides a unique opportunity for this study, given the public policy concerns about higher rates of family instability, especially among socioeconomically disadvantaged populations, and its implications for child wellbeing (Ventura and Bachrach 2000). The FFCWS collects detailed information on both the number and the type of family structure transitions from a child's birth to age 9. It also provides assessments of children's development at multiple times during early and middle childhood, periods that are critical to a variety of life course outcomes. Because these data include measures of children's cognitive achievement and externalizing and internalizing problem behaviors, we are able to investigate the possibility that family instability effects are outcome-specific.

Our analysis is based on 2,952 mother-child pairs. We exclude mothers who were lost to follow-up ($n = 1{,}791$), who lived less than half time with their focal child ($n = 102$), and who did not report their complete family structure history ($n = 53$).[3] For missing observations on covariates due to item-nonresponse, we use a multiple imputation (MI) procedure (Allison 2002). MI uses observed data to replace missing values with multiple imputed data and then obtains estimates averaged over these complete data with appropriate standard errors that take the uncertainty about sampling and imputation model into account. MI relies on weaker assumptions than do listwise deletion and other standard procedures for handling missing data (Little and Rubin 2002). Our MI procedure accounts for a rich array of variables that may be important factors for missing data patterns as well as for family instability and child development. We also include variables for family instability and child development during the MI procedure, but we exclude any missing observations on these variables in the analysis.[4] Our estimation is based on 20 imputed datasets created with the imputation using chained equations (ICE) option in Stata (Royston 2004). Although the study sample is slightly socioeconomically advantaged compared to the baseline sample, it is still based on a socioeconomically disadvantaged population. We then converted the study sample to person-year data for panel data analysis. Sample sizes vary by outcomes, with 6,525 person-years

[2]Fathers' response rates are based on mothers who completed the baseline interview.

[3]Mother-child pairs that were lost to follow-up include those who permanently dropped out of the survey and those who left the survey but rejoined later. We address the issue of sample attrition by incorporating censoring weights in our marginal structural models. To make the study samples consistent across models, we include mother-child pairs for analysis only if they are observed in all waves.

[4]To evaluate the sensitiveness of our MI strategy, we compare our estimates to those based on two alternative MI strategies. One set of estimates includes imputed cases on the treatment variable but excludes imputed cases on the outcome variables. The other set includes imputed cases on both the treatment and outcome variables. The results (available on request) show that while the findings are substantively similar across different MI strategies, the estimates based on the two alternative MI strategies are somewhat smaller in magnitude and have larger standard errors than the estimates based on our original strategy. We also note that our MI procedure utilizes all covariates as they are likely correlated with the missing data mechanisms (for recent research on missing data, see von Hippel 2007; Young and Johnson 2015).

for cognitive achievement[5] and 7,946 person-years for externalizing and internalizing behaviors.

Measures

Dependent variables—This study uses three key measures of child cognitive and socioemotional development, evaluated at Years 3, 5, and 9. Children's cognitive achievement is measured by the Peabody Picture Vocabulary Test-Revised (PPVT-R), which assesses the size and range of words that children understand. Two measures of children's socioemotional development are derived from the Child Behavioral Checklist (Achenbach and Rescorla 2000). Mothers responded to a series of items that pertain to their children's externalizing and internalizing problem behaviors. Each item consists of a three-point Likert scale on which mothers reported whether their children's behavior is "not true (0)," "sometimes or somewhat true (1)," or "often or very true (2)."

Externalizing problem behavior is measured by the sum of the aggressive and rule-breaking behavior subscales ($\alpha \approx .88$ across waves). The aggression subscale consists of items that ask about disobedience at home or at school, getting in fights, attacking people, screaming, and being usually loud. The rule-breaking subscale contains items that ask whether children hang around with others who get in trouble, cheat, prefer being with older children, run away from home, set fires, steal at or outside of home, swear, and vandalize.

Internalizing problem behavior is measured by the sum of the anxious/depressive and withdrawn behavior subscales ($\alpha \approx .82$ across waves). The anxious/depressive subscale consists of items that ask whether children fear they might think or do something bad, worry that they have to be perfect, complain no one loves them, feel guilty, are easily embarrassed, and worry in general. The withdrawn subscale contains items on being alone rather than with others, uninvolved in social activities, secretive, shy, underactive, and refusing to talk.

For cognitive achievement, a higher score represents a better outcome. For externalizing and internalizing behaviors, a higher score represents a worse outcome. The analysis standardizes all of these child developmental outcomes to have a mean of 0 and a standard deviation of 1, such that the effects of family instability are expressed in standard deviation units.

Family instability—We construct two measures of family instability: the number and the type of family structure transitions between survey waves. Both measures capture the immediate, acute impacts of family instability (Beck et al. 2010; Meadows et al. 2008). First, we measure the number of transitions in family structure (0, 1, and 2+) that a focal child's mother experienced between birth and age 3, between ages 3 and 5, and between ages 5 and 9. This measure is one of the most common approaches to gauging family instability. At each survey, mothers reported whether they were living with a partner, and if so, whether the current partner was the same person identified in the prior survey. If a mother experienced a coresidential exit or entrance between two surveys, we code this as one transition. If both occurred, we code it as two transitions. The Years 5 and 9 surveys

[5]We do not use scores on the cognitive achievement test conducted in Spanish because of their incompatibility with those in English.

asked mothers how many partners they had lived with since the prior surveys. We use this information to recode the number of family transitions between Years 3 and 5 and between Years 5 and 9 to account for transitions that were not picked up by mothers' reports on relationship status at each survey. This question was not asked at the Year 3 survey, so we may undercount family transitions before age 3. We treat the number of family transitions as a categorical measure because, as will be seen, associations between family instability and child outcomes are nonlinear, and only a small number of mothers experienced more than two transitions between surveys.

Our second measure of family instability focuses on the type of transition and is based on the distinction between living in a coresidential union (two-parent family) versus living in a single-mother family at the beginning of each period. It also accounts for multiple transitions between periods. This measure has six categories: (1) stable coresidential union; (2) moving out of a coresidential union into a single-mother family (one transition); (3) multiple transitions from a coresidential union; (4) stable single-mother family; (5) moving out of a single-mother family into a coresidential union (one transition); and (6) multiple transitions from a single-mother family. We measure these types of transitions between birth and age 3, between ages 3 and 5, and between ages 5 and 9.

Past research typically estimates family instability effects by comparing families experiencing particular types of family transitions with stable two-parent families. However, this specification may lead to comparisons between groups that are often not comparable. From a child's perspective, moving out of a two-parent family cannot occur if a child already lives in a single-mother family; likewise, moving into a two-parent family cannot occur if a child already lives in a two-parent family. To properly capture family instability effects, our analysis compares children who experience the exit of a father or father figure with children living in a stable two-parent family—(1), (2), and (3)—and children who experience the entry of a father or father figure with children living in a stable single-mother family—(4), (5), and (6).

Our measure of a coresidential union does not distinguish between biological and social fathers; nor does it distinguish between married and cohabiting parents. We collapse these categories because of small cell sizes for each of the disaggregated types of family transitions. For example, our preliminary analysis (see Table S1 in the online supplement) shows that only 47 and 22 children who were living with a biological father or in a married-parent family, respectively, experienced multiple transitions between ages 3 and 5. Also, there is no exit of a social father between birth and age 3 because family structure at birth is measured only with respect to a mother's relationship with a biological father. These aspects of our data likely lead to imprecise estimates and prevent us from estimating the effects of some types of family transitions. Nevertheless, we discuss results based on more fine-grained measures of family structure transitions in the Results section.

Covariates—Our analytic models include a rich array of covariates that have been shown to influence both family instability and child developmental outcomes, alongside a host of covariates that are typically treated as unobservable in prior research. Table 1 reports descriptive statistics of all covariates used in this study. Time-constant covariates consist of maternal, paternal, and child characteristics, measured at baseline. Maternal characteristics include age, race/ethnicity (black, Hispanic, white, and other), immigration status (1 if immigrant; 0 if otherwise), educational attainment (less than high school, high school or GED, some college, and college degree or more), age at first birth, cognitive ability (a subtest score of Wechsler's [1981] Adult Intelligence Scale-Revised), impulsivity (an abbreviated form of Dickman's [1990] dysfunctional impulsivity scale), and whether she lived with both parents at age 15 (1 if yes; 0 if no). Biological fathers' characteristics consist of age, mix-race couple, immigration status, educational attainment, employment status (not employed, working part-time, or working full-time), and incarceration status (1 if ever incarcerated; 0 if otherwise). Child characteristics include gender (1 if male; 0 if female), first-born status (1 if yes; 0 if no), and low birthweight status (1 if below 2,500 grams; 0 if otherwise).

We also construct a set of time-varying covariates that are known to be correlated with family structure transitions. Based on mothers' reports at baseline and ages 3 and 5, we measure poverty status (poverty, near poverty, or no poverty), union status (married to biological father, married to social father, cohabiting with biological father, cohabiting with social father, or single),[6] employment status (not employed, part-time, or full-time), living with a parent (child's grandparent) (1 if yes; 0 if no), number of children, depression status, domestic violence, religious attendance (from never to every day), alcohol/drug problems (1 if any; 0 if otherwise), and physical health status (poor, fair, good, very good, or excellent). Depression status (1 if yes; 0 if no) is based on the Composite International Diagnostic Interview-Short Form (CIDI-SF) (Kessler et al. 1998).[7] Domestic violence (1 if yes; 0 if no) is based on mothers' reports of any exposure to physical and emotional violence and coercive control by their spouse or partner. We also include children's age (in months) as a time-varying covariate to account for variation in the time of assessment of child developmental outcomes at each age.

ANALYSIS PLAN

This study uses three panel models to estimate the effects of family instability—the number and the type of family structure transitions—on each of the child developmental outcomes. Studies of family instability often compare children who have ever experienced a family transition with children who have never experienced a transition. In contrast, our approach examines period-specific instability rather than cumulative instability. Our reference group in each period is thus families that are stable during a specific period, rather than families that have been stable since the focal child's birth. This approach likely understates the effect of family instability on child outcomes, because some children in the reference group will have experienced instability in an earlier period.

We begin by estimating a random-effects model (REM) that takes the following form:

[6]Because union status at a point in time is correlated with the number of family transitions, we treat union status as a time-varying covariate in models that estimate the effects of the number of family transitions. Union status at baseline refers to a mother's relationship with her child's biological father.

[7]We use maternal depression measured at Year 1 as a baseline covariate because this measure is not available at birth.

$$\boldsymbol{Y}_{ti}=\beta_0+\beta_1 FI_{ti}+\beta_2 Year_{ti}+\beta_3 Year_{ti}^2+\beta_4 Month_{ti}+\beta_5 Month_{ti}^2+\boldsymbol{TC}_i\gamma+\boldsymbol{TV}_i\theta+u_{0i}+u_{1i}FI_{ti}+u_{2i}Year_{ti}+u_{3i}Year_{ti}^2+\varepsilon_{ti} \quad (1)$$

A vector of child developmental outcomes at time t for child i ($\boldsymbol{Y}$) is a function of a measure of family instability (FI), time ($Year$ and $Year^2$), child i's age in months at time t ($Month$ and $Month^2$), a vector of time-constant covariates ($\boldsymbol{TC}$), a vector of time-varying covariates ($\boldsymbol{TV}$), and random components (u's).[8] The model includes linear and quadratic terms for time to allow the time function to be nonlinear. The parameter estimate of interest is β_1, which estimates the association between family instability and the outcomes. The REM produces unbiased estimates of the effects of family instability on child outcomes under the assumption that, conditional on $\boldsymbol{TC}$ and $\boldsymbol{TV}$, FI is uncorrelated with the random effects and the idiosyncratic error (ε_{ti}). The REM has two problems. First, we cannot be certain that this assumption is correct. If the assumption does not hold, the parameter estimates of family instability will be biased. Second, even if the assumption holds and omitted variables bias is not a problem, the REM is insufficient because it does not allow us to properly account for measured variables that vary over time and that are correlated with family instability and the outcome of interest.

To deal with the first problem—bias due to confounding by unmeasured variables that are constant over time—we use a child fixed-effects model (FEM) to estimate the effects of family instability. The FEM takes the following form:

$$\boldsymbol{Y}_{ti}=\beta_0+\beta_1 FI_{ti}+\beta_2 Year_{ti}+\beta_3 Year_{ti}^2+\beta_4 Month_{ti}+\beta_5 Month_{ti}^2+\alpha_i+\varepsilon_{ti} \quad (2)$$

In this model, the child-specific intercept (α_i) denotes the deviation of each child's intercept from the mean intercept (β_0), representing all characteristics that are stable over time, whether they are observed or not. α_i is often estimated by including all of the indicators representing each child or demeaning both the outcome and explanatory variables by the child's overall mean. As Equation 2 shows, the FEM identifies parameter estimates by exploiting within-child variation in family structure transitions and child developmental outcomes.[9]

To deal with the second problem—bias due to the presence of measured covariates that vary over time—we estimate a marginal structural model (MSM) using inverse probability of treatment (IPT) weighting (Robins 1999; Robins, Hernán, and Brumback 2000).[10] To estimate the MSM, the IPT weighting first calculates the probability, p, that a child will have

[8]We also considered more complex REMs by interacting family instability with time and by allowing the coefficient of time to vary across time-constant covariates. These specifications produced similar results but made it difficult to compare with other panel models estimated in this study.

[9]The FEM can be inefficient if within-child variation is limited. Our inspection of the study sample shows this is not the case. For example, 13 percent of children living in a coresidential union from birth to age 3 experienced at least one transition by age 5, and 21 percent of children living in a coresidential union from ages 3 to 5 experienced at least one transition by age 9.

experienced family instability by a given time, conditional on prior history of family instability and observed time-constant and time-varying covariates. Then it weights each child by the inverse of her conditional probability. Children in the treated group (those who have experienced a family structure change at time *t*) are given a weight of $1/p$, thereby assigning lower weights to children with higher probabilities and higher weights to children with lower probabilities. Children in the comparison group (those who did not experience the treatment at time *t*) are given a weight of $1/(1 - p)$, thereby assigning higher weights to children with higher probabilities and lower weights to those with lower probabilities. The MSM thus generates a pseudo-population in which family transitions are sequentially independent of prior observed covariates.

Given that we adjust for time-varying covariates (e.g., poverty status) that are potentially confounded with family structure transitions in our pseudo-population, it is no longer necessary to condition on time-varying covariates that may operate as mediators in the models predicting child developmental outcomes. By minimizing the problems of confounding by and over-controlling for observed time-varying covariates, the MSM overcomes an important drawback of conventional panel models, namely, their inability to adequately handle the reciprocal relationship between family structure transition and its time-varying covariates.

Let $FI_t = fi$ denote child *i*'s actual treatment status at time *t*. For time-varying covariates, we use overbars to denote covariate history up to time *t*: $\overline{\boldsymbol{TV}}_t = \{\boldsymbol{TV}_0, \boldsymbol{TV}_1, \ldots, \boldsymbol{TV}_t\}$. Because a small number of observations with extreme weights (outliers) may distort the estimation process (Hernán, Brumback, and Robins 2002), we compute stabilized IPT weights to increase efficiency:

$$\mathrm{IPTW}_{ti} = \prod_{t=0}^{T} \frac{Pr(FI_t = f_i | \overline{FI}_{t-1}, \boldsymbol{TC})}{\Pr(FI_t = fi | \overline{FI}_{t-1}, \boldsymbol{TC}, \overline{\boldsymbol{TV}}_{t-1})} \quad (3)$$

Π is the product operator; the denominator is the probability that child *i* received the actual treatment of family instability at time *t*, conditional on prior family instability history, and time-constant and time-varying covariates at time $t - 1$; and the numerator is the probability that child *i* received the actual treatment of family instability at time *t*, conditional on prior family instability history, and time-constant covariates.[11] We compute IPT weights by fitting a series of pooled logit regression models in which we contrast each treatment of family structure transitions with the corresponding comparison groups.[12] For example, children experiencing two or more transitions are contrasted with those experiencing no transition.

[10]The MSM shares with propensity score matching the common goal of randomizing treatment assignment based on observed covariates. However, when the treatment of interest is time-varying, the MSM has a clear advantage over the propensity score matching method. In matching, children in the comparison group at time *t* are discarded if they are not matched to children in the treated group at time *t*, even though they can be in either the treated group or the matched comparison group at time $t + 1$. The matching method is therefore likely to produce a severely truncated subset of data, leading to inefficient estimation. The MSM does not pose this problem as it retains the entire dataset.

[11]Weights are truncated at the 1st and 99th percentiles to avert disproportionate influence from outlying observations (Cole and Hernán 2008). Our inspection (available on request) indicates that the MSMs are well-behaved, as the weight distribution is centered around values close to 1, has small variance, and is only slightly skewed to the right.

[12]Further analysis (not shown) suggests these models do not violate the independence of irrelevant alternatives assumption.

Children moving out of a two-parent family are contrasted with those living in a stable two-parent family.

As with any panel data, sample attrition is inevitable in our data. Nonrandom attrition, in turn, may yield biased results. We address this issue by constructing weights for time-varying exposure to censoring (Robins et al. 2000). Let $L_t = 1$ if child i was lost to follow-up by time t and $L_t = 0$ if otherwise. $\bar{L}_{t-1} = 0$ denotes that child i was not lost to follow-up by time $t - 1$. The stabilized censoring weights are given as follows:

$$\mathrm{CW}_{ti}=\prod_{t=0}^{T}\frac{Pr(L_t=0|\overline{L}_{t-1}=0,\overline{FI}_{t-1},\boldsymbol{TC})}{Pr(L_t=0|\overline{L}_{t-1}=0,\overline{FI}_{t-1},\boldsymbol{TC},\overline{\boldsymbol{TV}}_{t-1})} \quad (4)$$

The MSM estimates family instability effects with the product of the IPT and censoring weights (IPTW_{ti} x CW_{ti}), fitting a random-effects model.[13] We control for baseline covariates as they enter into both the numerator and denominator of the stabilized weights. Throughout the analysis, we compute robust standard errors to correct for within-individual correlation.

RESULTS

Descriptive Results

We begin by describing the distribution of children by number and type of family transitions (Table 2). Among children experiencing any family structure transition, the majority experience only one. In a supplemental analysis (not shown), we found that among children experiencing two or more transitions, the majority experience only two (86 percent between birth and age 3, 92 percent between ages 3 and 5, and 75 percent between ages 5 and 9). During early to middle childhood, family transitions occur fairly equally across family type.

Table 2 also shows differences in mean levels of age-specific child outcomes. At age 3, children who experience at least one family transition between birth and age 3 have lower levels of cognitive achievement and higher levels of externalizing behavior than children who experience no transition. Levels of internalizing behavior are significantly higher for children who experience only one family transition. Comparisons by types of family structure transitions indicate that these differences are more common among children born into two-parent families than among children born into single-mother families. Compared to children in stable coresidential unions, children who experience the exit of a father suffer developmentally from such transitions. However, compared to children in stable single-mother families, children who experience any transitions do not significantly differ. Overall, these patterns are repeated for the periods between ages 3 and 5 and between ages 5 and 9.

[13]The MSM assumes no unobserved time-constant confounding, conditional on observed covariates. To evaluate the robustness of our model to unobserved heterogeneity, we conducted a falsification test based on the premise that future family transitions cannot affect current child outcomes. In this test, we included family transitions between ages 5 and 9 as a covariate in the model predicting child outcomes at ages 3 and 5. Results (available on request) indicate that although MSM estimates are sensitive to bias in some cases, unobserved heterogeneity does not substantively alter the main findings reported here. We thank an anonymous reviewer for suggesting this test.

Multivariate Results

Next we present estimates of the effects of family instability on children's developmental outcomes during early to middle childhood (Tables 3 to 5). In each table, panel A reports results from models that examine the number of family structure transitions, and panel B reports results from models that examine the type of family transition. We estimate four models: a random-effects model (REM) with time-constant covariates only, a REM with time-constant and time-varying covariates, a child fixed-effects model (FEM), and a marginal structural model (MSM). Throughout these analyses, both REMs yield almost identical estimates.

Table 3 reports results for cognitive achievement. In panel A, the REM estimates (columns 1 and 2) show that multiple family structure transitions (2+) have a significant effect, lowering children's PPVT score by .06 standard deviations (SDs). The MSM estimate (column 4) is not significant but similar in magnitude, suggesting that confounding by observed time-varying covariates is not a concern. The FEM estimate (column 3) is smaller than the other estimates and statistically insignificant, indicating that the REM and MSM estimates may be biased by unobserved time-constant covariates.

Next we examine whether the type of family instability matters (panel B). Here, the REM estimates indicate that family instability—both moving out of a coresidential union and experiencing multiple family transitions—significantly reduces children's PPVT scores (column 2: −.068 SDs, $p < .01$ and −.102 SDs, $p < .01$, respectively). The MSM yields similar estimates. Although the FEM estimates are statistically insignificant, they are similar in size and direction to the REM and MSM estimates, at least for a move out of a coresidential union. The FEM estimate for multiple transitions from a coresidential union is half the size of the REM and MSM estimates, suggesting the estimates based on the REM and MSM may be biased by unobserved time-constant differences. Finally, estimates at the bottom of Table 3 indicate that having a father or father figure move into the household does *not* reduce children's cognitive achievement. Rather, the signs on the coefficients are positive and the FEM estimate (.078 SDs) is significant. The fact that the FEM estimate is twice as large as the REM estimates suggests that unobserved time-constant characteristics associated with a move in may be masking the benefits of this type of transition.

Table 4 reports results for children's externalizing behavior. All the models in panel A indicate that family instability increases children's externalizing behavior. The estimates are consistent, suggesting the REM is not biased by time-constant or time-varying confounding. The REM and MSM estimates for two or more transitions are similar in magnitude to those for one transition, but the FEM estimate is smaller. The estimates displayed in panel B, however, suggest that the effect of multiple transitions is large and significant when the type of transition is taken into account. Children who begin the period in a two-parent family and who experience multiple transitions exhibit more externalizing behaviors than do their peers who remain in a stable two-parent family. The REM, FEM, and MSM estimate that the effect is about .15 SDs ($p < .001$), 11 SDs ($p < .05$), and .12 SDs ($p < .05$), respectively.

Table 5 presents estimates for children's internalizing behavior. These results (panel A) show that experiencing one family structure transition leads to more internalizing behavior, even in the presence of various forms of selection bias (REM: about .06 SDs, $p < .01$; FEM: .061 SDs, $p < .01$; MSM: .049 SDs, $p < .05$). In panel B, although only the REM coefficients are statistically significant, the estimates for a move out of a coresidential union are comparable across models. The strongest effects are seen for multiple moves out of coresidential unions (.093 to .102 SDs). Although these effects are not precisely estimated in the FEM and MSM, they are almost identical in magnitude to those estimated from the REM. Surprisingly, children who experience multiple family transitions from a single-mother family have a lower level of internalizing behavior than do children who remain in stable single-mother families. This effect is sensitive to selection bias, however, due to unobserved time-constant confounding as indicated by the FEM.

In summary, the analyses reported in Tables 3, 4, and 5 show that family structure instability has a negative effect on children's cognitive and socioemotional development during early to middle childhood, even after accounting for different types of selection bias. These results also suggest that focusing only on the quantity of family structure transitions may lead to misleading conclusions. In the case of cognitive achievement, the effect of family instability depends on the type of change. Whereas a move out of the household reduces a child's cognitive achievement, a move into the household has a positive although insignificant effect. In the case of behavior problems, the effect of a family structure transition is generally negative regardless of type of move, but typically larger and more significant for moves out of a two-parent family.

We also estimate models that distinguish between biological and social fathers and between married and cohabiting two-parent families. We find negative effects of transitions out of living with a biological father (see Table S2 in the online supplement). The FEM estimates indicate that multiple transitions out of a two-biological-parent family increase children's externalizing and internalizing behaviors by .179 SDs ($p < .001$) and .129 SDs ($p < .05$), respectively. Our results also indicate that transitions out of a married-parent family increase behavior problems (see Table S3 in the online supplement). According to the FEM estimates, multiple transitions out of a married-parent family increase children's externalizing and internalizing behaviors by .233 SDs ($p < .01$) and .189 SDs ($p < .05$), respectively. In short, the estimates obtained from the disaggregated family transition measures are more substantial than those obtained from the current measure. For children living in a single-mother family, we do not find consistent patterns using the disaggregated measures. As noted earlier, small cell sizes may produce imprecise estimates and prevent us from estimating the effects of some types of family transitions. With this caution in mind, the results suggest that the family instability effects reported in our main analysis are driven by transitions from living with a biological father and from living in a married-parent family.

Subgroup Analysis

Findings from our main analysis suggest that (1) transitions from a two-parent family lower children's PPVT score (Table 3); (2) transitions from a two-parent family and, to a lesser degree, a transition into a two-parent family increase children's externalizing behavior (Table 4); and (3) transitions from a two-parent family and transition into a two-parent family increase children's internalizing behavior, although the effects are not always significant (Table 5). To investigate heterogeneous responses to family instability by race/ethnicity and by child gender, we replicate the previous specifications separately for each subgroup.[14] Because the purpose of our subgroup analysis is to examine whether family instability effects differ across racial/ethnic and gender groups, we focus on more robust estimates from child fixed-effects and marginal structural models.

Prior studies have found that family instability effects are more salient for whites than for other racial/ethnic groups (Fomby and Cherlin 2007; Wu and Thomson 2001). Table 6 reports our results, which suggest that racial/ethnic differences depend on the outcome examined. For cognitive achievement (panel A), effects of family instability are larger for blacks than they are for other groups, although the differences in the coefficients across racial/ethnic groups are not statistically significant. For externalizing behavior (panel B), the effects of family instability are larger for Hispanic children, especially children who experience the entrance of a father figure into the household (a multiple transition from a two-parent family typically involves a move out of a two-parent family followed by a move into another two-parent family). There is also suggestive evidence that moving out of a two-parent family increases white children's externalizing behavior. Finally, for internalizing behavior (panel C), the effect of moving out of a two-parent family is greater for white children, and the effect of moving into a two-parent family is greater for Hispanic children.

Table 7 reports results for gender differences. As discussed earlier, prior studies suggest that family instability has more deleterious consequences for boys than for girls, especially for behavior problems (Cavanagh et al. 2008; Cooper et al. 2011). Our findings are consistent with prior work on behavior problems but not cognitive outcomes. As panel A shows, the exit of a parent from the household has a larger negative effect on girls' cognitive achievement, although the gender difference is not statistically significant. In contrast, the detrimental effect of family instability on externalizing behavior is more pronounced for boys than for girls, especially for boys who experience multiple transitions from a two-parent family (panel B). An unexpected finding is that multiple family transitions from a single-mother family appear to lower externalizing behavior among girls. Results for internalizing behavior are similar to those for externalizing behavior (panel C). Multiple family transitions from a two-parent family, and single transitions from a single-mother family, have a more negative effect on boys than on girls. Interestingly, both types of transitions involve the entrance of a father figure into the household: a move out and in for children in a two-parent family, and a move in for children in a single-mother family. Once again, there is some evidence that multiple transitions from a single-mother family reduce internalizing behavior among girls, although this effect is much smaller and insignificant in the FEM.

In summary, whereas the effect of family instability on cognitive achievement is greater for black children compared to white or Hispanic children, and for girls compared to boys, the effect of family instability on socioemotional development is greater for Hispanic and white children and for boys. These effects also differ by the type of family transition for each population subgroup. Effects of a transition out of a coresidential union are more

[14]The analysis excludes other racial/ethnic groups due to their small sample size.

pronounced for black and white children than for Hispanic children, whereas effects of transitions into a coresidential union are more salient for Hispanic children. These findings suggest that inferences about family instability effects based on the general population are likely incomplete, as they may mask heterogeneity across population subgroups.

DISCUSSION

Family changes over the past half century have created fundamental shifts in children's experiences of family life (Bumpass and Lu 2000; U.S. Census Bureau 2006; Ventura 2009). Driven by higher rates of divorce/separation, remarriage/repartnering, nonmarital childbearing, and cohabitation, children in the United States are growing up in increasingly dynamic family structures. The present study adds to the literature on family instability by assessing the role of family structure transitions in shaping children's developmental outcomes during early to middle childhood. Using data from the FFCWS, we extend prior research in several important ways. First, we examine the number and the type of family structure transitions to determine whether it is the quantity or the type of transition that matters most for children's development. Second, we provide greater analytic leverage than prior research by accounting for time-varying confounding by observed covariates, an overlooked source of selection bias, as well as time-constant confounding by unobserved covariates. Third, we address population heterogeneity by examining how family instability effects differ by race/ethnicity and by children's gender.

On the whole, we find that distinguishing between moves out of a two-parent family and moves into a two-parent family is important, casting doubt on the claim that all types of instability are equally harmful for children. Generally speaking, transitions out of a two-parent family are more harmful to children than transitions into a two-parent family. Furthermore, we find a good deal of population heterogeneity in the effects of family transitions on child development. For cognitive achievement, the impact of transitions out of two-parent families is stronger and more negative for black children and girls, although the differences across groups are not statistically significant. For socioemotional development, the impact of transitions out of two-parent families is stronger and more negative for white children, whereas the impact of transitions into two-parent families is stronger and more negative for Hispanic children. These differences are statistically significant.

So how large are these effects? To address this question, we use estimates from the marginal structural models to compare the magnitude of the family instability effect with the magnitude of the effect of maternal education and poverty status, both of which are universally acknowledged to be consequential for children's development. Figure 1 shows that the effect of family structure instability on cognitive achievement is about one-third the size of the effect of having a mother with high school education (versus college or more) and about one half the size of the effect of being born into a poor household (versus no poverty). In the case of children's cognitive development, parental socioeconomic status is clearly more important than family structure instability. For socioemotional development, however, the story is different. Family structure instability has a larger effect on children's externalizing behavior than does maternal education or poverty status, and a comparable effect on children's internalizing behavior. These findings are consistent with a growing body of research that finds family structure instability affects children's future success primarily by reducing their socioemotional skills or mental health (Heckman 2007; McLanahan et al. 2013).

As noted earlier, we focus on family structure transitions from two-parent or single-mother families. But two-parent families are a diverse group who differ by paternal relationship with children (biological or social father) and marital status (marriage or cohabitation). To examine diversity in family instability, we conducted additional analyses using more fine-grained measures of family structure transitions. These results suggest that the negative effects of transitions from a two-parent family are likely driven by transitions involving a biological father and married parents, rather than transitions involving a social father and cohabitating parents. However, given imprecise estimation due to small cell sizes, we interpret these results as suggestive. More research using larger panel data is needed to substantiate the findings of this study.

Our analysis also reveals that the effects of family instability on children's development may differ by population subgroups. With respect to race/ethnicity, previous studies suggest that family instability effects may be more pronounced for whites than for racial/ethnic minority groups, owing to differential processes of selection into family instability. Racial/ethnic minority groups, for example, are more likely than whites to experience nonmarital childbearing and family instability, encounter multiple stressors arising from structural disadvantage, and utilize kinship networks to absorb the adverse effects of family instability (Amato and Keith 1991; Fomby and Cherlin 2007; McLoyd et al. 2000). Our results for white children's socioemotional development are consistent with prior research. We also find, however, that transitions out of a two-parent family have more adverse effects on black children's cognitive achievement, whereas transitions into a two-parent family have more adverse effects on Hispanic children's socioemotional development. One reason why our findings may differ from those of past studies is that most studies do not simultaneously distinguish by type of family structure change and examine multiple child outcomes.

The results for Hispanic children were surprising. In supplemental analyses (not shown), we found that the negative effect of a move into a two-parent family on the socioemotional development of Hispanic children was most pronounced among children of immigrant mothers who moved in with a social father (mostly cohabitation). Children in this group also had more siblings than those in stable single-mother families, suggesting that higher levels of behavior problems may be driven in part by new births. Because the Hispanic families in these data are heterogeneous in terms of country of origin and immigration status, these findings should be interpreted with caution and subjected to further investigation.

With respect to children's gender, prior research predicts gendered responses to family instability, contending that changes in father figures are more disruptive for boys than for girls (Allison and Furstenberg 1989; Cooper et al. 2011). Consistent with this view, we find that the adverse effects of family instability on children's socioemotional development are more salient for boys than for girls, especially multiple family transitions originating from a two-parent family. We also find some evidence that family structure transitions may reduce the internalizing behavior of girls, which is inconsistent with prior research. Note, however, that most studies that report a negative effect are based on adolescent girls (McLanahan et al. 2013). It is possible that the negative consequences of family instability do not appear among girls until they reach adolescence. Taken together, our findings underscore the importance of population heterogeneity for better understanding the association between family instability and child development.

Our study has several limitations. First, although we estimate family instability effects using child fixed-effects and marginal structural models to address selection bias, neither of these models can simultaneously account for unmeasured time-constant and time-varying covariates. Thus we cannot rule out the possibility that unmeasured confounders that change over time are responsible for the association between family structure transitions and child development. In addition, our approach is designed to estimate the effects of repeated treatments on repeated outcomes. Alternatively, one could take latent class modeling approaches (e.g., latent trajectory and latent transition analyses) to identify longitudinal profiles of family instability (Collins and Lanza 2010; Wagmiller et al. 2006). By taking a holistic view, these approaches account for the interdependence of family transitions over time. We consider our approach and longitudinal latent class modeling approaches as complementary to one another in enhancing our understanding of family instability effects.

A second limitation of our study is that sample size prevents us from examining the effects of instability in dating partnerships (Beck et al. 2010; Cooper et al. 2011). The fact that single mothers frequently engage in unstable dating relationships may account for the fact that family transitions are mostly insignificant for children living in single-mother families. Also, given our focus on children who are living with their mother, the coverage of family transitions is incomplete (e.g., children who are living with only their grandparents or in foster care are not included). Although data limitations remain a challenge, a more complete conceptualization and measurement of family instability is an important agenda for future research on this topic. Third, information on the number of coresidential partners between survey years is not available from birth to age 3, which means we must rely on mothers' relationship status at two points in time to assess family structure transitions during early childhood. This limitation undoubtedly leads us to underestimate the prevalence of family structure instability during this period. Finally, the FFCWS data are representative of families living in large cities (populations of 200,000 or more). We thus caution that the results reported here may not represent family instability effects for the U.S. population as a whole. Still, our findings are germane to children whose exposure to family structure instability is relatively high.

In closing, our findings underscore the fact that family instability effects are (1) contingent on the type of family structure transitions and (2) subject to considerable heterogeneity by race/ethnicity and gender. They also indicate that family instability has important consequences for children's cognitive and socioemotional development, even after various forms of selection bias are taken into account. We therefore propose that more attention be given to the type of family transition to which children are exposed and to population heterogeneity in children's responses to instability. Such a nuanced understanding of family instability will benefit policymakers as they try to effectively allocate private and public resources to fostering family stability and reducing inequalities in child development.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

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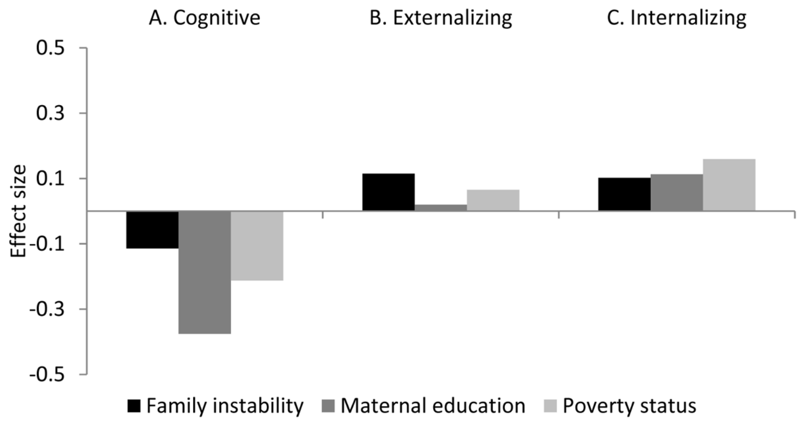

Figure 1.

Comparison of Effect Sizes of Family Instability, Maternal Education, and Poverty Status

Note: Results are from MSMs using the full sample. The effects of family instability indicate experiencing multiple transitions from a coresidential union as compared to a stable coresidential union. The effects of maternal education indicate high school versus college or more. The effects of poverty status indicate poverty versus no poverty.

Table 1

Descriptive Statistics of Control Variables

	Baseline		Age 3		Age 5	
	Mean or %	**(SD)**	**Mean or %**	**(SD)**	**Mean or %**	**(SD)**
Maternal Characteristics						
Age (in year)	25.25	(6.04)				
Race/Ethnicity						
Black	50.24%					
Hispanic	24.20%					
White	22.04%					
Other	3.52%					
Immigrant	13.46%					
Education						
Less than high school	30.44%					
High school or GED	31.47%					
Some college	26.10%					
College or more	11.99%					
Poverty Status						
Poverty	33.88%		40.11%		39.09%	
Near poverty	26.36%		25.54%		26.69%	
No poverty	39.77%		34.35%		34.21%	
Living Arrangement						
Married-biological	25.61%		32.98%		32.87%	
Married-social			1.59%		3.90%	
Cohab-biological	35.09%		19.10%		12.94%	
Cohab-social			7.12%		10.95%	
Single	39.30%		39.21%		39.34%	
Employment						
Not employed	6.87%		40.88%		39.01%	
Part-time	29.91%		15.54%		16.62%	
Full-time	63.22%		43.58%		44.37%	

	Baseline		Age 3		Age 5	
	Mean or %	**(SD)**	**Mean or %**	**(SD)**	**Mean or %**	**(SD)**
Age at First Birth	21.73	(5.29)				
Cognitive Ability	6.87	(2.67)				
Impulsivity	6.06	(3.61)				
Living with Parent at Age 15	42.37%					
Living with Parent	26.43%		14.96%		11.97%	
Number of Children	1.25	(1.29)	2.32	(1.31)	2.52	(1.32)
Depression	15.83%		20.41%		16.89%	
Domestic Violence	13.19%		47.80%		35.61%	
Religious Attendance	2.10	(1.37)	2.66	(1.34)	2.69	(1.36)
Alcohol/Drug	13.09%		7.48%		7.22%	
Health	2.93	(0.94)	2.74	(1.03)	2.66	(1.02)
Father Characteristics						
Age (in year)	27.79	(7.24)				
Mixed-Race Couple	14.46%					
Immigrant	14.74%					
Education						
Less than high school	31.13%					
High school or GED	35.30%					
Some college	22.47%					
College or more	11.10%					
Employment						
Not employed	2.06%					
Part-time	11.66%					
Full-time	86.28%					
Ever Incarcerated	26.90%					
Child Characteristics						
Age (in month)			35.55	(2.40)	61.52	(2.64)
Male	52.44%					
First Born	39.17%					
Low Birthweight	10.03%					

Note: N = 2,952.

Table 2

Child Developmental Outcomes, by Family Instability

	Number of Family Transitions			Type of Family Transitions					
				Coresidential			Single		
	0	1	2+	Stable	Move out	Multiple	Stable	Move in	Multiple
Age 3									
PPVT	.135 (1.002)	−.098[a] (.925)	−.154[a] (.930)	.268 (1.018)	−.035[b] (.963)	−.101[b] (.905)	−.094 (.930)	−.161 (.881)	−.197 (.953)
Externalizing	−.077 (.970)	.041[a] (1.012)	.147[a] (1.003)	−.177 (.897)	−.023[b] (1.009)	.196[b] (.962)	.112 (1.071)	.113 (1.009)	.095 (1.028)
Internalizing	−.081 (.980)	.045[a] (1.013)	.011 (.974)	−.196 (.908)	.029[b] (0.999)	.031[b] (.883)	.137 (1.074)	.056 (1.029)	−.003 (1.060)
Percent	61.61%	26.30%	12.09%	41.50%	13.18%	6.08%	20.11%	13.11%	6.01%
Age 5									
PPVT	.225 (.931)	−.097[a] (.899)	−.165[a] (.939)	.347 (.915)	−.090[b] (.892)	−.389[b] (1.029)	.007 (.920)	−.102 (.904)	−.056 (.872)
Externalizing	−.085 (.895)	.084[a] (.916)	.97[a] (.911)	−.165 (.840)	.061[b] (.970)	.093[b] (.922)	.075 (.977)	.105 (.861)	.100 (.905)
Internalizing	.037 (.986)	.139[a] (.938)	.041 (.913)	.018 (.986)	.136 (.929)	.166 (1.128)	.073 (.985)	.141 (.945)	−.021 (.777)
Percent	69.53%	22.14%	8.34%	47.08%	11.27%	2.51%	22.40%	10.90%	5.84%
Age 9									
PPVT	.106 (.982)	−.104[a] (.900)	−.201[a] (.805)	.300 (.978)	−.189[b] (.819)	−.200[b] (.806)	−.201 (.907)	−.010[c] (.971)	−.208 (.796)
Externalizing	−.065 (.916)	.158[a] (1.119)	.226[a] (1.138)	−.131 (.889)	.155[b] (.990)	.273[b] (1.264)	.048 (.960)	.158 (1.225)	.180 (.906)
Internalizing	−.044 (.912)	.039 (1.054)	.142[a] (1.215)	−.051 (.919)	.040 (.969)	.216[b] (1.386)	−.033 (.908)	.039 (1.135)	.027 (.849)
Percent	67.23%	24.69%	8.08%	43.26%	12.64%	4.89%	23.96%	12.06%	3.19%

Note: N = 2,952. Means and standard deviations are presented.

[a]Indicates that the mean differs statistically from no family transitions.

[b]Indicates that the mean differs statistically from a stable coresidential family.

[c]Indicates that the mean differs statistically from a stable single-mother family.

Table 3

Effects of Family Structure Transitions on Cognitive Achievement

	REM w/o TV	REM w/TV	FEM	MSM
	(1)	(2)	(3)	(4)
A. *Number of Family Structure Transitions*				
0 (ref.)				
1	−.013 (.024)	−.007 (.024)	.010 (.028)	−.014 (.025)
2+	−.063* (.034)	−.062* (.035)	−.018 (.039)	−.058 (.039)
B. *Type of Family Structure Transitions*				
Stable coresidential union (ref.)				
Move out	−.064* (.034)	−.068** (.034)	−.067 (.042)	−.076** (.036)
Multiple transitions from coresidential union	−.105** (.049)	−.102** (.049)	−.050 (.057)	−.114** (.058)
Stable single motherhood (ref.)				
Move in	.040 (.034)	.047 (.034)	.078* (.041)	.000 (.037)
Multiple transitions from single motherhood	−.022 (.048)	−.028 (.048)	.009 (.055)	−.044 (.050)

Note: N = 6,525 person-years. Robust standard errors are in parentheses. All models include survey year, its square term, child's age, and its square term. REM and MSM also control for observed time-constant covariates.

TV = time-varying covariates.

* $p < .05$;

** $p < .01$;

*** $p < .001$ (one-tailed tests).

Table 4

Effects of Family Structure Transitions on Externalizing Behavior

	REM w/o TV	REM w/TV	FEM	MSM
	(1)	(2)	(3)	(4)
A. *Number of Family Structure Transitions*				
0 (ref.)				
1	.057** (.024)	.053** (.024)	.055** (.027)	.051* (.027)
2+	.063* (.035)	.061* (.035)	.025 (.039)	.046 (.039)
B. *Type of Family Structure Transitions*				
Stable coresidential union (ref.)				
Move out	.065* (.034)	.062* (.034)	.029 (.041)	.051 (.035)
Multiple transitions from coresidential union	.154*** (.050)	.149*** (.050)	.111* (.057)	.115* (.060)
Stable single motherhood (ref.)				
Move in	.051 (.034)	.051 (.035)	.080** (.040)	.050 (.040)
Multiple transitions from single motherhood	−.021 (.048)	−.020 (.048)	−.050 (.053)	−.038 (.051)

Note: N = 7,946 person-years. Robust standard errors are in parentheses. All models include survey year, its square term, child's age, and its square term. REM and MSM also control for observed time-constant covariates.

TV = time-varying covariates.

* $p < .05$;

** $p < .01$;

*** $p < .001$ (one-tailed tests).

Table 5

Effects of Family Structure Transitions on Internalizing Behavior

	REM w/o TV (1)	REM w/TV (2)	FEM (3)	MSM (4)
A. *Number of Family Structure Transitions*				
0 (ref.)				
1	.058** (.025)	.059** (.025)	.061** (.029)	.049* (.027)
2+	.019 (.037)	.019 (.037)	.020 (.042)	.011 (.043)
B. *Type of Family Structure Transitions*				
Stable coresidential union (ref.)				
Move out	.070** (.035)	.065* (.035)	.067 (.044)	.052 (.036)
Multiple transitions from coresidential union	.098* (.053)	.093* (.053)	.098 (.061)	.102 (.065)
Stable single motherhood (ref.)				
Move in	.049 (.036)	.049 (.037)	.056 (.044)	.038 (.040)
Multiple transitions from single motherhood	−.052 (.051)	−.049 (.051)	−.048 (.058)	−.087* (.052)

Note: N = 7,946 person-years. Robust standard errors are in parentheses. All models include survey year, its square term, child's age, and its square term. REM and MSM also control for observed time-constant covariates.

TV = time-varying covariates.

* $p < .05$;

** $p < .01$;

*** $p < .001$ (one-tailed tests).

Table 6

Effects of Family Structure Transitions on Child Developmental Outcomes, by Race/Ethnicity

	Black		Hispanic		White	
	FEM	MSM	FEM	MSM	FEM	MSM
A. *Cognitive Achievement*						
Stable coresidential (ref.)						
Move out	−.101*	−.096**	−.045	−.078	−.034	−.059
Multiple transitions from coresidential union	−.081	−.158**	−.042	−.016	−.018	−.081
Stable single (ref.)						
Move in	.077	−.008	.009	.007	.212*	.037
Multiple transitions from single motherhood	−.035	−.055	.067	−.082	.089	−.105
N (person-years)	3,474		1,469		1,356	
B. *Externalizing Behavior*						
Stable coresidential (ref.)						
Move out	.051	.029	−.051	.048	.058	.182**a,b
Multiple transitions from coresidential union	.104	.093	.271**c	.262*a,c	.004	.097
Stable single (ref.)						
Move in	.026	−.019	.238**a,c	.188*a	.026	.110
Multiple transitions from single motherhood	−.038	−.060	−.015	−.029	−.220	.027
N (person-years)	4,027		1,878		1,750	
C. *Internalizing Behavior*						
Stable coresidential (ref.)						
Move out	.028	−.025	.016	.078	.192**a,b	.229***a,b
Multiple transitions from coresidential union	.021	.027	.217	.168	.140	.198
Stable single (ref.)						
Move in	.018	−.015	.185*a	.177*	.034	.098
Multiple transitions from single motherhood	−.048	−.078	−.058	−.044	−.155	−.241
N (person-years)	4,027		1,878		1,750	

Note: All models include survey year, its square term, child's age, and its square term. MSM also controls for observed time-constant covariates. The coefficients of these covariates are not shown to conserve space.

[a] Indicates that the coefficient differs statistically from blacks.

[b] Indicates that the coefficient differs statistically from Hispanics.

[c] Indicates that the coefficient differs statistically from whites.

* $p < .05$;

** $p < .01$;

*** $p < .001$ (one-tailed tests).

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Table 7

Effects of Family Structure Transitions on Child Developmental Outcomes, by Child Gender

	Boy		Girl	
	FEM	**MSM**	**FEM**	**MSM**
A. *Cognitive Achievement*				
Stable coresidential (ref.)				
Move out	−.023	−.044	−.101*	−.106**
Multiple transitions from coresidential union	−.072	−.084	−.023	−.164**
Stable single (ref.)				
Move in	.071	−.037	.081	.044
Multiple transitions from single motherhood	.087	.004	−.076	−.090
N (person-years)	3,399		3,126	
B. *Externalizing Behavior*				
Stable coresidential (ref.)				
Move out	.063	.066	.000	.036
Multiple transitions from coresidential union	.152*	.127	.056[a]	.092
Stable single (ref.)				
Move in	.107*	.056	.051	.042
Multiple transitions from single motherhood	.042	.094	−.147**	−.188***[a]
N (person-years)	4,166		3,780	
C. *Internalizing behavior*				
Stable coresidential (ref.)				
Move out	.050	.014	.083	.077
Multiple transitions from coresidential union	.190**	.177*	−.024[a]	.005[a]
Stable single (ref.)				
Move in	.124**	.054	−.017[a]	.014
Multiple transitions from single motherhood	.010	.018	−.110	−.222***[a]
N (person-years)	4,166		3,780	

Note: All models include survey year, its square term, child's age, and its square term. MSM also controls for observed time-constant covariates. The coefficients of these covariates are not shown to conserve space.

[a]Indicates that the coefficient differs statistically from boys.

* $p < .05$;

** $p < .01$;

*** $p < .001$ (one-tailed tests).