Job Displacement among Single Mothers: Effects on Children's Outcomes in Young Adulthood[1]

Jennie E. Brand and Juli Simon Thomas
University of California, Los Angeles

Given the recent era of economic upheaval, studying the effects of job displacement has seldom been so timely and consequential. Despite a large literature associating displacement with worker well-being, relatively few studies focus on the effects of parental displacement on child well-being, and fewer still focus on implications for children of single-parent households. Moreover, notwithstanding a large literature on the relationship between single motherhood and children's outcomes, research on intergenerational effects of involuntary employment separations among single mothers is limited. Using 30 years of nationally representative panel data and propensity score matching methods, the authors find significant negative effects of job displacement among single mothers on children's educational attainment and social-psychological well-being in young adulthood. Effects are concentrated among older children and children whose mothers had a low likelihood of displacement, suggesting an important role for social stigma and relative deprivation in the effects of socioeconomic shocks on child well-being.

Given the deep and broad recent economic downturn, studying the effect of job displacement on families has seldom been so timely and consequential. One in seven children in the United States under age 18 (i.e., 10.5 million children) had an unemployed parent in 2010, a number nearly

[1] This article is dedicated to the memory of our colleague Suzanne Bianchi, who was the first to provide useful comments and suggestions on the original draft of this manuscript and whose work is an inspiration to those who study work and family. We also thank Greg Duncan, Till von Wachter, James Raymo, Judith Seltzer, Megan Sweeney, Yana Kucheva, and the *AJS* reviewers for useful comments and suggestions. Versions of this

0002-9602/2014/11904-0002$10.00

double that of those who had an unemployed parent before the downturn (Lovell and Isaacs 2010). An even larger number of children had a parent displaced from a job, a form of involuntary job loss that occurs as a result of the operating decisions of employers, that is, when firms downsize, restructure, close plants, or relocate, as opposed to job separation that occurs when workers are fired or leave jobs voluntarily. Yet worker displacement is not limited to economic downturns. Over the last several decades, economic reorganization in the United States has been characterized by waves of job displacement, increasingly widespread job insecurity, and the perceived disappearance of a lifetime job with a single employer for growing segments of the workforce (Levy 1995; Wetzel 1995; Farley 1996; Kalleberg 2000, 2009; Farber 2010). Prior research suggests that job displacement is associated with subsequent unemployment, long-term earnings losses, and lower job quality; declines in physical and mental well-being; new patterns of interaction and involvement with family and peers; and loss of psychosocial assets and personal assessment in relation to societal norms and unemployment stigmatization (Jahoda, Lazarsfeld, and Zeisel [1933] 1971; Pearlin et al. 1981; Jahoda 1982; Podgursky and Swaim 1987; Newman 1988; Ruhm 1991; Seitchik 1991; Jacobson, LaLonde, and Sullivan 1993; Fallick 1996; Kletzer 1998; Chan and Stevens 2001; Charles and Stephens 2004; Farber 2005; Brand 2006; Brand and Burgard 2008; Cha and Morgan 2010; Couch and Placzek 2010; von Wachter 2010; Catalano et al. 2011).

Notwithstanding this body of research on the effects of displacement on workers, we know relatively less about the effects of parental displacement on children and, thus, about the intergenerational impact of a consequential household event. A few recent studies find significant negative effects of parental job loss on children's academic achievement, socioeconomic status, and social-psychological well-being (e.g., Kalil and Ziol-Guest 2005, 2008; Oreopoulos, Page, and Stevens 2008; Page, Stevens, and Lindo 2009; John-

article were presented at the Population Association of America 2011 Annual Meeting, the American Sociological Association 2012 Annual Meeting, the Advanced Quantitative Methods in Education Research seminar series at UCLA, the California Center for Population Research (CCPR) and the Family Working Group at UCLA, the Department of Sociology at the University of Arizona, and the Institute for Research on Poverty 2013 Summer Research Workshop at the University of Wisconsin—Madison. This research made use of facilities and resources at CCPR, UCLA, which receives core support from the Eunice Kennedy Shriver National Institute of Child Health and Human Development (grant R24HD041022), and received support from National Institutes of Health (grant 1 R01 HD07460301A1). This research was conducted with restricted access to Bureau of Labor Statistics data. The ideas expressed herein are those of the authors. Direct all correspondence to Jennie E. Brand, Department of Sociology, University of California, Los Angeles, 264 Haines Hall, Los Angeles, California 90095. E-mail: brand@soc.ucla.edu

son, Kalil, and Dunifon 2012). However, these studies largely treat fathers as the essential component in the displacement equation; indeed, little work has examined the consequences of job displacement among mothers, and even fewer studies have focused on single mothers. As a large proportion of children in the United States will spend some part of their childhood raised by a single mother, and single mother families are a vulnerable population for which job displacement may be acutely taxing, we attend to effects on the children of displaced single mothers. While there is a large literature on the socioeconomic and social-psychological outcomes of children who grow up with single mothers (see McLanahan and Percheski [2008] and Crosnoe and Cavanagh [2010] for recent reviews), much of this work treats single mothers' labor market experience as a static state: women are either employed or unemployed, each with its associated trials for raising children. Although some recent research takes trajectories of labor force experiences among single mothers more seriously, this work largely treats those career trajectories as voluntary. Such studies hence face challenging problems of selection bias. Less subject to selection concerns, the study of job displacement provides a unique opportunity to assess how socioeconomic hardship among single mothers affects the well-being of children. Displacement not only provides a sounder basis for studying the intergenerational impact of socioeconomic circumstances, but a substantial number of parents separate from jobs as a result of economic conditions rather than voluntary choices, and we should understand the potential intergenerational consequences of such events.

In this article, we present evidence as to the effects of job displacement among single mothers on children's subsequent educational and social-psychological outcomes in young adulthood. Using data from the National Longitudinal Survey of Youth (NLSY) and from the National Longitudinal Survey's Child-Mother File (NLSCM), we match children whose single mothers were and were not displaced from their jobs but had the same propensity to be displaced on the basis of a set of observed covariates. Matching is an intuitive approach for adjusting estimates of causal effects for confounding due to covariates that may be correlated with the causal variable and the outcome variable; the approach has been used in several prior studies of job displacement (e.g., Brand 2004, 2006; Brand and Burgard 2008; Couch and Placzek 2010). We also study variation in maternal displacement effects, including heterogeneous displacement effects by the observed propensity for maternal displacement (Xie, Brand, and Jann 2012; Brand and Simon Thomas 2013), multistate displacement effects by contexts with more and less concentrated unemployment, and time-varying displacement effects by children's age when maternal displacement occurs (Brand and Xie 2007).

BACKGROUND AND SIGNIFICANCE

Parental Job Displacement Effects on Children

As job displacement has substantial, long-term effects on workers' socioeconomic status, physical and psychological well-being, and social involvement, we reasonably expect these consequences to negatively affect children of displaced workers. A decrease in parental economic resources may restrict the ability to purchase goods critical for child development, such as schooling, housing, food, and cognitively enriched learning environments (Conger and Elder 1994; Duncan and Brooks-Gunn 1997; Kalil and Ziol-Guest 2008). The association between job loss and home loss via mortgage default and foreclosure, landlord foreclosure, or inability to pay rent and reliance on public assistance suggests additional mechanisms by which parental displacement can affect the well-being of children (Antel 1992). Parental socioeconomic shocks and downward mobility can also dampen children's confidence in the value of education and work.

Moreover, to the extent that displaced parents model despondency and despair, they may foster psychological distress among children. McLoyd and Wilson (1990) note, "deterioration in the parent's psychological functioning in the context of economic loss or poverty may become a communicable social phenomenon to the extent that the child imitates the symptomatic affect and behavior of the parent" (p. 53). Displaced parents' decreased psychological well-being can also inhibit emotional warmth and incite erratic or punitive parenting practices leading to poor adjustment in children (Elder 1974, 1998; Elder, Nguyen, and Caspi 1985; Kessler, Turner, and House 1989; McLoyd 1990, 1998; McLoyd and Wilson 1990; McLeod and Shanahan 1993; Conger and Elder 1994; McLoyd et al. 1994; Conger, Conger, and Elder 1997). Children are particularly vulnerable to maternal depression, with social-psychological and behavioral effects persisting through adulthood (Elder 1998; Meadows, McLanahan, and Brooks-Gunn 2007; Turney 2011). Although periods of unemployment potentially allow parents more time to spend with children, the stress associated with financial and employment uncertainty and the time needed to devote to job search instead seemingly reduce the quantity and quality of time spent with children (Lareau 2011; Brooks-Gunn and Markman 2005). As families lacking social capital and collective efficacy have children with lower educational and socioeconomic attainment, displaced parents' geographic mobility and withdrawal from social life may negatively affect children's outcomes (McLanahan 1983; Coleman 1988, 1990; Haveman, Wolfe, and Spaulding 1991; Astone and McLanahan 1994; Sandefur and Laumann 1998; Furstenberg et al. 1999; Sampson, Morenoff, and Earls 1999; Leventhal and Brooks-Gunn 2000).[2]

[2] It may be that the change in neighborhood characteristics typically associated with geographic mobility is the cause of declines in child well-being (Jackson and Mare 2009).

Given the many reasons to expect a link between parental displacement and children's outcomes,[3] and the current timeliness of studying the consequences of economic contractions, a literature is emerging on the effects of parental displacement on children. Most of this literature has focused on short-term educational and social-psychological outcomes, including lower self-esteem and higher likelihood of grade repetition, dropout, and suspension or expulsion from school (Kalil and Ziol-Guest 2005, 2008; Stevens and Schaller 2011; Johnson et al. 2012).[4] Yet as displacement is associated with enduring outcomes for displaced workers themselves, it may likewise exert long-term effects on children. One study finds a significant negative effect on children's educational attainment at age 21 (Kalil and Wightman 2011); another study finds parental displacement is associated with the lower income of children in adulthood (Page et al. 2009). Studies on parental displacement in Canada and Norway find negative effects on children's educational attainment and positive effects on social assistance in adulthood (Oreopoulos et al. 2008; Coelli 2011; Rege, Telle, and Votruba 2011). These studies, however, largely emphasize the deleterious effects of fathers' loss of financial standing in the family among married couple households. Those studies examining differences between paternal and maternal displacement effects among married couples find significant effects of paternal but not maternal displacement (Kalil and Ziol-Guest 2008; Rege et al. 2011); they generally hypothesize that maternal displacement is not as detrimental to children's outcomes as paternal because of greater psychological consequences associated with employment and income loss among fathers who are largely expected to maintain the role of primary provider. However, these studies do not explicitly examine the effects of displacement among single mothers.

Single Parenthood, Maternal Employment, and Children's Outcomes

The employment experiences of single mothers are important for a significant population of children in the United States, as about half will spend a portion of their childhood raised by a single parent (Ellwood and Jencks 2004; McLanahan and Percheski 2008; Brown 2010), about 84% of custodial single parents are mothers, and about 80% of single mothers are employed (U.S. Census Bureau 2009). Single working mothers are more

[3] Job displacement is also associated with divorce (Charles and Stephens 2004). However, as we focus on single mothers at the time of displacement in this article, we do not review the links between displacement and divorce or between divorce and children's outcomes.

[4] Another strand of work shows that community-level job losses affect the achievement test scores of children, possibly the result of both direct effects on children whose parents lost jobs and indirect peer and teacher effects (Ananat, Gassman-Pines, and Gibson-Davis 2011).

likely to be black and Hispanic, be low income, have lower earnings and lower earnings potential, live in poverty, and endure job displacement and spells of unemployment relative to married working mothers (Harris 1993; Edin and Lein 1997*a*, 1997*b*; Attewell 1999; Cancian and Meyer 2004; Ellwood and Jencks 2004; McLanahan 2004, 2009; Edin and Kefalas 2005; McLanahan and Percheski 2008; Brown 2010). Children raised by single mothers have on average worse educational, socioeconomic, and social-psychological outcomes, ostensibly the result of economic deprivation, reliance on public assistance, psychological strain, inconsistent parenting style, low social control, and residence in disadvantaged communities (Dornbusch et al. 1985; McLanahan 1985, 2004; Amato 1987; Amato and Keith 1991; Astone and McLanahan 1994; McLanahan and Sandefur 1994; Seltzer 1994; Pebley and Sastry 2004; McLanahan and Percheski 2008; Brown 2010; Crosnoe and Cavanagh 2010).[5] Attention to the impact of labor market work among single mothers has increased, especially as welfare reform has meant that a larger proportion of single mothers is in the labor force and faces even greater consequences to work interruption (Danziger et al. 2002; Crosnoe and Cavanagh 2010). Past research has focused on the constraints single mothers face to serving the dual role of nurturer and provider (Conger and Elder 1994; McLoyd et al. 1994; Edin and Lein 1997*a*, 1997*b*; Hao and Brinton 1997; Scott et al. 2001; Harknett 2006). This discourse and research on employment among single mothers largely treats labor force participation as a static and voluntary state. Yoshikawa et al. (2006) comment: "What is missing from this research literature on the effects of work on children? Astonishingly, almost none of the studies on maternal work and children's development examine the impact of changes in maternal work and its conditions on children" (p. 12).

There are, of course, noteworthy exceptions to this leaning, that is, studies that have examined the relationship between employment trajectories, job instability, and children's outcomes among single-parent families (Kalil, Duniform, and Danziger 2001; Chase-Lansdale et al. 2003; Kalil and Ziol-Guest 2005; Yoshikawa et al. 2006; Hill et al. 2011; Stevens and Schaller 2011). Yet while this work captures the complexity of employment trajectories among single mothers, it does not generally differentiate between voluntary and involuntary job separation and is thus faced with significant issues of selection bias. Several of these studies employ research designs to address concerns with selection, such as instrumental variables and quasi-

[5] There is considerable family status complexity and instability within the "single" mother population and thus heterogeneity in the social and emotional support, as well as strain, mothers and children experience (Yoshikawa, Weisner, and Lowe 2006; Brown 2010). Nevertheless, variation in child well-being outside of two-parent married families is comparatively low (Brown 2010).

experimental designs, but in so doing limit the generalizability of their findings to specific subpopulations. By contrast, displacement events are relatively exogenous to individual characteristics (i.e., relative to employment trajectories and voluntary job transitions). Stevens and Schaller (2011), for example, find little difference in the effects of parental displacement on children's outcomes using alternative estimation strategies, including fixed effects, leading them to assert that effects of displacement are not driven by unobserved heterogeneity. Indeed, several recent studies of parental job displacement have used the displacement event as a purer estimate of the effects of parental income shocks on children's outcomes (Oreopoulos et al. 2008; Page et al. 2009; Stevens and Schaller 2011).[6] In addition to reducing concerns over selection, estimating the effects of involuntary employment separations should be a focus of study for understanding the potential intergenerational consequences of the recent recession.

Variation in Parental Job Displacement Effects on Children

It is implausible to assume that children respond identically to parental displacement. We first hypothesize that displacement effects on children vary by parents' likelihood of displacement, which is influenced by worker characteristics and economic conditions. Displacement that occurs during recessions, in which many workers are laid off, is associated with greater economic losses than displacement that occurs during economic expansions (Fallick 1996; von Wachter 2010; Couch, Jolly, and Placzek 2011; Davis and von Wachter 2011). However, contexts of widespread unemployment lessen the internalization of blame and social stigma associated with job loss (Miller and Hoppe 1994; Clark 2003, 2010; Charles and Stephens 2004; Brand et al. 2008). That is, displaced workers may benefit from a "social norm effect": as aggregate unemployment increases, one's own unemployment represents a smaller deviation from the social norm (Clark 2010), and thus displacement effects on social-psychological well-

[6] Although studying displacement reduces concerns over selection bias, we do not mean to say that we fully mitigate endogeneity issues. Employers may lay off workers they deem to be less productive, motivated, or committed to the job, or for a host of other idiosyncratic reasons. One additional strategy to deal with endogeneity concerns is to separate layoffs from plant closings, where the effects of plant closings provide a less biased estimate of displacement. However, we are not convinced this is the appropriate interpretation of such differences, and we do not employ this as a strategy to deal with selection in our main analyses. Job losses due to layoffs and those due to plant closings are also potentially different treatment conditions. Layoffs may produce larger effects because of lower postdisplacement productivity signals (irrespective of actual worker characteristics; Gibbons and Katz 1991); the loss of a job from larger, higher-wage employment establishments (Krashinsky 2002); or greater postdisplacement internalization, stigmatization, and fewer similarly strained workers to offer a network of support (Miller and Hoppe 1994; Brand, Levy, and Gallo 2008).

being are often less in contexts of mass layoffs. Likewise, while more disadvantaged workers, for whom displacement is more likely, may be more vulnerable to financial shocks, such economic adversity is a comparatively normative experience; by contrast, job displacement and socioeconomic decline may instigate an acute sense of deprivation among more advantaged families whose peers tend to be likewise advantaged and for whom displacement is a relative shock. Prior parental displacement research has examined variation in effects on socioeconomic attainment by worker characteristics, finding that effects are concentrated among disadvantaged families (Oreopoulos et al. 2008; Kalil and Wightman 2011; Stevens and Schaller 2011). However, this work does not consider how effects vary within a sample of single-parent households, effects on children's social-psychological well-being, or how effects vary by the likelihood of displacement.

We further hypothesize that child age when parental job displacement occurs interacts with the effects of displacement, an issue that has received little attention to date. As early childhood is important for cognitive development, health, and social and emotional development (Duncan and Brooks-Gunn 1997; Duncan et al. 1998; Shonkoff and Phillips 2000; Dahl and Lochner 2005; Duncan, Ziol-Guest, and Kalil 2010), it may be a period especially sensitive to parental displacement and associated economic adversity. Conversely, periods of unemployment allow mothers more time to spend with children, potentially deflating the deleterious effects of economic hardship among preschool-age children. Moreover, young children are less subject to the social-psychological effects of relative socioeconomic status declines, whereas older children are more attuned to social stigma and relative status (Conger et al. 1997).[7] Economic adversity is also critical to adolescent well-being, for their extracurricular activities and social networks (Mistry et al. 2001), and in their educational decision-making process.[8]

DATA, DESCRIPTIVE STATISTICS, AND METHOD

NLSY and NLSCM

We use data from the National Longitudinal Survey of Youth (NLSY; Bureau of Labor Statistics and U.S. Department of Labor 2008) and the Na-

[7] Some studies found no effects of low-income single mothers' employment when children are young on children's socioeconomic outcomes, and negative effects if employment is marked by poor working conditions, but positive effects of maternal employment when children are adolescents (see Crosnoe and Cavanagh [2010] for a review).

[8] In our analyses, we focus on variation in effects by the treatment condition rather than by the sample condition. Another potentially interesting hypothesis is that maternal displacement effects within a "single" mother population vary by family composition, e.g., by partnership status, characteristics of partners, or marital status at the time of birth. We explore several such models in results not shown and do not find systematic differences; however, in future work we continue to explore these questions.

tional Longitudinal Survey's Child-Mother File (NLSCM; Bureau of Labor Statistics, U.S. Department of Labor, and National Institute for Child Health and Human Development 2008). The NLSY is a nationally representative sample of 12,686 respondents who were 14–22 years old when first surveyed in 1979; these individuals were interviewed annually through 1994 and biennially thereafter. In 1986, the National Longitudinal Survey began a separate survey of the children of NLSY women, the NLSCM. Data have been collected every two years since 1986, with new sections added in 1994 as children entered young adulthood. As of 2010, the 6,283 NLSY women had given birth to about 11,500 children. The NLSY is an underused resource for the study of job displacement (Kletzer and Fairlie 2003). In addition to extensive information on individual characteristics, NLSY data distinguish between layoffs and firings, which stands in contrast to many surveys used to study displacement.[9] It also has a short recall period for reporting job losses for a panel survey, from one (in 1984–94) to two (in 1996–2010) years, an important component for retrospective reports of unemployment (Duncan and Mathiowetz 1988).

We define displacement as termination from the mother's main job between each interview period as a result of layoff or plant closing; in other words, we consider mothers who quit jobs or were fired as nondisplaced.[10] As the NLSY grouped layoffs with temporary job endings before 1984, we do not use pre-1984 data on displacement. These years also correspond to early labor market activity, marked by job churning, and are not the primary focus of most displacement studies. We merge data on women from the NLSY with data on children from the NLSCM (n = 4,931 mothers and n = 11,504 children, where children become our unit of analysis) and reconstruct displacement indicators from year (1984–2010) to child age (0–17). We restrict our sample to children whose mothers were not continuously married and living with a spouse between 1984 and 2010 and for whom we have data on mothers' job separation (n = 3,109 mothers and n = 6,751 children) and to children who were at least 19 years old in 2010 (n = 2,834 mothers and n = 5,697 children), as all of our outcomes pertain to children who are at least 19 years old. Single motherhood is not a static state. Displaced mothers were single when they were displaced, as were nondisplaced mothers at some period during their children's childhood, but not necessarily continuously single. Our primary indicator of dis-

[9] Measurement error is nonetheless possible if respondents report they were laid off when they were in fact fired, a problem common to all survey data attempting to make this distinction.

[10] We restrict displacement to the Current Population Survey (CPS) job, or the main job respondents held in each spell. Our restriction, like that employed in definitions of displacement by the Bureau of Labor Statistics, is motivated to establish worker attachment to jobs lost and not to categorize multiple overlapping jobs lost as displacements.

placement captures maternal job loss that occurs anytime during childhood (ages 0–17). Alternative displacement indicators are described below.

Covariates used to estimate the propensity for maternal displacement are described in table 1. Most of theses measures are straightforward. In 1980, 94% of NLSY respondents (the mothers) were administered the Armed Services Vocational Aptitude Battery (ASVAB), a battery of 10 intelligence tests measuring knowledge and skill in areas such as mathematics and language. We first residualize each of the ASVAB tests on age at the time of the test separately by race and ethnicity. Residuals were standardized to mean 0 and variance 1. We construct a scale of the standardized residuals ($\alpha = .92$) with mean 0, standard deviation 0.8, and a range of -3 to 3. We incorporate two measures of personality characteristics (or "noncognitive" skills). First, we include the Rotter Locus of Control Scale collected in 1979, designed to measure the extent to which individuals believe they have control over their lives through self-motivation or self-determination (internal control) as opposed to the extent that the environment (chance, fate, luck) controls their lives (external control). The scale ranges from 4 to 16, where lower values indicate greater internal control. Although the internal consistency of the scale is low for the NLSY cohort ($\alpha = .36$), the scale correlates well with self-esteem, education, and social class. Second, we construct a measure of delinquent activity on the basis of 16 questions collected in 1980 asking whether the respondent engaged in activity such as stealing, gambling, fighting, or doing drugs. We use several additional time-varying predisplacement characteristics of mothers: number of children, married or partnered,[11] years of education, region of residence (northeast and north central or south and west), employer tenure (measured in weeks), and whether mothers were manufacturing or trade workers. For the full sample, covariates are measured at children's birth (i.e., predisplacement interval), but we also construct each of these time-varying measures for children at ages 6 and 12, that is, as baseline measures for each time-varying displacement interval.

In order to assess the impact of maternal displacement during childhood on both educational attainment and social-psychological well-being of young adults, we explore five outcomes: high school completion by age 19, college attendance by age 21, college completion by age 25, and depressive symptoms at ages 20–24 and 25–29. Depressive symptoms in young adulthood were collected in 1994–2010 and measured with a seven-item Center for Epidemiological Studies Depression Scale (CESD; Radloff 1977) that asked respondents whether they never/rarely, sometimes, occasionally, or

[11] Although women are single when they are displaced, they may have different relationship statuses throughout the course of their children's lives, and we therefore include married or partnered status as a baseline covariate.

TABLE 1
DESCRIPTIVE STATISTICS: NLSY WOMEN (1979–2010) AND NLSCM CHILDREN (1986–2010)

	MOTHER NOT DISPLACED CHILD 0–17		MOTHER DISPLACED CHILD 0–17		
	Mean/ Proportion	SD	Mean/ Proportion	SD	T-TESTS
Maternal time-invariant covariates:					
Black (0/1)	.248	. . .	.389	. . .	***
Hispanic (0/1)	.101	. . .	.067	. . .	
(Grand)mother's years of education	10.429	2.699	10.577	2.399	+
(Grand)parents' intact family age 14 (0/1)	.643	. . .	.648	. . .	*
(Grand)parents' southern residence (0/1)	.386	. . .	.400	. . .	**
Mental ability (ASVAB; −3 to 3)	−.222	.730	−.359	.661	***
Rotter locus of control scale (6–14)	9.127	2.380	9.176	2.312	*
Delinquency (0/1)	.750	. . .	.843	. . .	**
Maternal covariates, child age 0:					
Age	21.315	2.816	22.128	3.097	**
Number of children	1.751	1.152	1.706	1.060	
Marital or cohabiting partner (0/1)	.676	. . .	.519	. . .	***
High school completion (0/1)	.737	. . .	.776	. . .	
College completion (0/1)	.028	. . .	.005	. . .	***
Region (0/1)	.430	. . .	.351	. . .	
Employer tenure (weeks)	12.099	19.081	10.447	17.299	**
Full-time employment (0/1)	.997	. . .	.986	. . .	
Manufacturing worker (0/1)	.126	. . .	.118	. . .	**
Trade worker (0/1)	.157	. . .	.164	. . .	
Maternal covariates, child age 6:					
Age	27.315	2.816	28.128	3.097	**
Number of children	2.451	1.206	2.395	.939	
Marital or cohabiting partner (0/1)	.606	. . .	.409	. . .	***
High school completion (0/1)	.776	. . .	.792	. . .	
College completion (0/1)	.039	. . .	.005	. . .	***
Region (0/1)	.413	. . .	.368	. . .	
Employer tenure (weeks)	21.514	31.264	17.980	26.842	***
Full-time employment (0/1)	.996	. . .	.994	. . .	
Manufacturing worker (0/1)	.124	. . .	.199	. . .	***
Trade worker (0/1)	.175	. . .	.231	. . .	
Maternal covariates, child age 12:					
Age	33.315	2.816	34.128	3.097	**
Number of children	2.756	1.245	2.644	1.036	
Marital or cohabiting partner (0/1)	.599	. . .	.399	. . .	***
High school completion (0/1)	.826	. . .	.824	. . .	
College completion (0/1)	.054	. . .	.029	. . .	***
Region (0/1)	.404	. . .	.365	. . .	
Employer tenure (weeks)	36.838	47.026	24.581	31.872	***
Full-time employment (0/1)	.822	. . .	.834	. . .	+
Manufacturing worker (0/1)	.119	. . .	.147	. . .	***
Trade worker (0/1)	.175	. . .	.251	. . .	***

TABLE 1 (*Continued*)

	Mother Not Displaced Child 0–17		Mother Displaced Child 0–17		
	Mean/ Proportion	SD	Mean/ Proportion	SD	T-Tests
Child covariate:					
Age in 2010	27.348	2.377	27.092	2.331	
Child outcomes in young adulthood:					
High school completion (0/1)	.760	. . .	.743	. . .	*
Some college attendance (0/1)	.427	. . .	.350	. . .	***
College completion (0/1)	.169	. . .	.149	. . .	**
CESD ages 20–24 (0–1)	.231	.177	.298	.225	*
CESD ages 25–29 (0–1)	.208	.181	.254	.238	**
Weighted sample proportion	.86		.14		

Note.—Sample restricted to children whose mothers were not continuously married and living with their spouse and for whom we have data on job separation, and who were at least 19 years old in 2010 ($n = 5{,}697$). Statistics weighted with an NLSY panel weight. For high school completion by age 19, $n = 4{,}412$; for college attendance by age 22, $n = 3{,}993$; for college completion by age 25, $n = 2{,}817$; for depressive symptoms ages 20–24, $n = 3{,}897$; and for depressive symptoms ages 25–29, $n = 3{,}192$. T-tests correspond to differences between children whose mothers were not and were displaced. Sample sizes differ according to missing values and different age thresholds for each dependent variable.

$^{+}P < .10$.

* $P < .05$.

** $P < .01$.

*** $P < .001$.

most/all of the time (1) had poor appetite, (2) had trouble keeping their mind on tasks, (3) were depressed, (4) felt that everything took extra effort, (5) had restless sleep, (6) were sad, and (7) could not get going. We perform factor analyses for each set of items in 1994–2010 and find that the seven items load onto a single factor in each year. We create scales for each year that range from 0 to 1; alphas range from 0.66 to 0.74. We then construct two measures that correspond to children ages 20–24 and 25–29, where we use the most recent assessment for children older than the lower-bound age.

Estimating Maternal Displacement Effects

For unit i, the effect of maternal displacement is defined as the difference between the two potential outcomes in the treated and the untreated states ($d = 1,0$; Rubin 1974; Imbens 2004; Heckman 2005; Morgan and Winship 2007):

$$\delta_i = y_i^{d=1} - y_i^{d=0}. \quad (1)$$

That is, we ask whether children whose mothers were displaced from jobs have different outcomes than they otherwise would have had if their moth-

ers had not been displaced. While less governed by selection than voluntary job separation, displacement is nonetheless conditioned by maternal characteristics that are also associated with levels of subsequent child outcomes. If unobserved maternal characteristics affect decisions to work at firms that are prone to instability and these characteristics are also correlated with parenting and outcomes of children, the effects of maternal displacement on children will be biased. Or, if firms make idiosyncratic layoff decisions, it may be that relatively less productive workers (e.g., workers with lower levels of motivation and ability) are also parents of children who have worse outcomes. The key to our identification strategy is the assumption that displacement is an exogenous event that is not correlated with such unobserved factors that could affect children's outcomes. Although we condition on a rich set of observed characteristics, including factors like cognitive ability and personality characteristics, we can never be sure that this assumption is fully satisfied.

We report a series of estimates of maternal displacement, beginning with simple bivariate associations, or unmatched mean differences, to provide a baseline estimate of differences in outcomes between children whose mothers were and were not displaced. We then report kernel-matching estimates in which children whose mothers were and were not displaced are matched according to their propensity for maternal displacement (Rosenbaum and Rubin 1983, 1984; Rubin 1997). We estimate the propensity score with a logit regression of the following form:

$$P_i = p(d_i = 1|X) = \log\frac{d_i}{1 - d_i} = \left(\sum_{k=0}^{K} \beta_k X_{ik}\right). \tag{2}$$

Differences between the unmatched and the matched estimates suggest selection into displacement by observed covariates. We then examine both the average treatment effect on the treated (TT),

$$E(\delta|d = 1, P) = E(y^{d=1} - y^{d=0}|d = 1, P), \tag{3}$$

and the average treatment effect on the untreated (TUT),

$$E(\delta|d = 0, P) = E(y^{d=1} - y^{d=0}|d = 0, P). \tag{4}$$

Differences between the TT and TUT suggest heterogeneity in treatment effects, an issue with substantive significance to which we return below. Appendix A provides more details regarding our matching estimators.[12]

[12] There are at least two advantages to using matching over conventional regression models. First, in contrast to regression, we make no functional form assumption for the relationship between the covariates and outcomes using matching. Second, covariate imbalance (also called "common support") is a focal concern in matching routines, while

Estimating Variation in Maternal Displacement Effects

Heterogeneous treatment effects.—We first consider variation in effects of maternal displacement on children by the observed likelihood of having a mother who is displaced from a job on the basis of a range of maternal characteristics. Our stratification-multilevel (SM) method to estimate heterogeneous treatment effects (Jann, Brand, and Xie 2010*a*; Xie et al. 2012) involves the following steps: (1) estimate propensity scores for each child for the probability of maternal displacement given a set of observed covariates ($P(d = 1 \mid X)$), (2) construct balanced propensity score strata such that there are no significant differences in the average values of covariates and the propensity scores between the children whose mothers were and were not displaced, (3) estimate propensity score stratum-specific maternal displacement effects (level 1 effect estimates), and (4) evaluate trends across strata using variance-weighted least squares regressions of the stratum-specific displacement effects obtained in step 3 on strata rank (level 2 effect estimates). We estimate our level 2 model by

$$\delta_s = \hat{\hat{\delta}}_1 + \gamma S + \epsilon_s, \tag{5}$$

where level 1 slopes (δ_s) are regressed on propensity score rank indexed by S, $\hat{\hat{\delta}}_1$ represents the level 2 intercept (i.e., the predicted value of the effect of maternal displacement for the lowest propensity stratum), and γ represents the level 2 slope (i.e., the change in the effect of maternal displacement on children's outcomes with each one-unit change to a higher propensity score stratum). Our objective is to look for a systematic pattern of treatment effects across strata. The assumption is that we consider all units within strata, treated and untreated, as homogeneous for estimating treatment effects. Although the assumption of within-stratum homogeneity is unlikely to be true, it is less stringent than the full-sample homogeneity assumption. This method has been used in empirical research on the effects of education (Brand 2010; Brand and Xie 2010; Brand and Davis 2011; Musick, Brand, and Davis 2012) and market processes in China (Xie and Wu 2005) but not for the effects of job transitions.

We also test sensitivity to the parametric and stratum-specific homogeneity assumptions imposed in SM, using a newly developed matching-smoothing (MS) method of estimating heterogeneous treatment effects (Xie

imbalance between treated and untreated cases goes undetected all too often in regression analyses. We do not use fixed effects models because the young adult outcomes we explore are not repeated before and after the displacement interval. We do not use sibling fixed effects models because the siblings share the treatment condition (i.e., siblings jointly experience maternal displacement, at least over their childhood). But even had siblings not shared the treatment, we question whether the fixed family effect in such a study would induce an endogenous selection bias.

et al. 2012) that involves the following steps: (1) estimate propensity scores for each unit, (2) match treated units to untreated units with a matching algorithm, (3) plot the observed differences between treated and untreated units against a continuous representation of the propensity score, and (4) use a nonparametric model (local polynomial with degree 1 and bandwidth 0.1) to smooth the variation in matched differences to obtain the pattern of treatment effect heterogeneity as a function of the propensity score.

Multistate treatment effects.—Second, we study variation in effects of maternal displacement on children's outcomes by considering multistate treatment conditions, where we separate displacement that occurred in contexts of more or less concentrated unemployment. That is, we examine separately the effects for children whose mothers were displaced during recessionary (i.e., 1990–92, 2001–2, 2008–10) and nonrecessionary periods (i.e., 1984–89, 1993–2000, 2003–7). Additionally, using private geocode data from the NLSY, we separate children whose mothers were displaced in labor market areas (LMAs; i.e., metropolitan or micropolitan areas for the computed balance of the state unemployment rate) with high and "nonhigh" unemployment, where high unemployment is defined as at least 9% unemployed.[13] The untreated groups in both cases include children whose mothers were not displaced. Use of alternative indicators also provides a measure of robustness to results, suggesting that context affects maternal displacement effects on children. Both of these multistate indicators test the degree to which maternal displacement effects on child well-being are more sensitive to economic adversity, which is generally greater in high-unemployment contexts, or social stigma and relative deprivation (i.e., the evaluation of relative standing vis-à-vis reference peers; Merton and Kitt 1950; Davis 1959), which are generally greater in lower-unemployment contexts. However, the macroeconomic context may involve less social salience for individual experiences than the regional context if reference comparisons are situated at more proximate levels of aggregation (Clark 2003).

Time-varying treatment effects.—Third, to study variation in effects of maternal displacement by children's age when displacement occurs, we construct three age-specific displacement intervals: 0–5 (early childhood years), 6–11 (middle childhood years), and 12–17 (adolescent years). Within each interval, displaced mothers must be single at the time of the event, and nondisplaced mothers must have been single at some point during the interval; both displaced and nondisplaced mothers could be married at other times. To address the counterfactual complexity involved in studying time-varying treatments, we adopt the conceptual framework of Brand and Xie

[13] Results are robust to alternative thresholds (i.e., 8% and 10%) for defining high unemployment.

(2007). Simplifying by restricting to nonrepeatable events, in this case initial maternal displacement, treatment can occur in period $d = t$, where $t = \{\text{ages } 0\text{–}5,\ 6\text{–}11,\ 12\text{–}17\}$. Analogous to an event history setup, children at risk for experiencing maternal displacement at time interval t have not experienced the event up to the baseline of t. The reference children include all those who have not experienced a maternal displacement up through time t and those who do and do not experience maternal displacement at any time subsequent to t. We thus remain agnostic about future events, treating future paths as unknown to us as they are to the individuals living them.[14] Figure 1 is a flowchart depicting our framework for estimating time-varying maternal displacement effects. We begin with all children, and those children's mothers can either be displaced when children are 0–5 or not displaced; those whose mothers are nondisplaced by age 5 can have mothers either displaced at 6–11 or not displaced; those whose mothers are nondisplaced by age 11 can have mothers either displaced at 12–17 or not displaced. Each transition is associated with a marginal probability weight $p(\cdot)$ of being treated or $q(\cdot)$ of not being treated in period t. We define the treatment effect at t on an outcome measured at T as

$$\delta_i^{t*} = y_i^{d=t} - y_i^{*d>t}, \tag{6}$$

where $y_i^{d=t}$ is the value of the outcome for children whose mothers were displaced in period $d = t$ and $y_i^{*d>t}$ is the value of the outcome for the same unit had that unit not been treated up until t. We define $y_i^{*d>t}$ as

$$y_i^{*d>t} = \sum_{h=1}^{T} w_{ih} y_i^{d=h} + w_{T+} y_i^{d>t}, \tag{7}$$

where w's are weights, we denote children whose mothers were not displaced in any of the observed periods by $d > T$, and we apply the following normalization constraint:

$$\sum_{h=1}^{T} w_{ih} + w_{T+} = 1. \tag{8}$$

We match children by the time-invariant covariates and time-varying covariates that correspond to child age before each treatment interval (i.e., ages 0, 6, and 12).

[14] For example, consider a child whose mother lost her job when the child was 8 years old, i.e., during the age interval 6–11. The reference child is one whose mother had not lost a job up through age 11 but whose mother may or may not have lost a job when the child was ages 12–17.

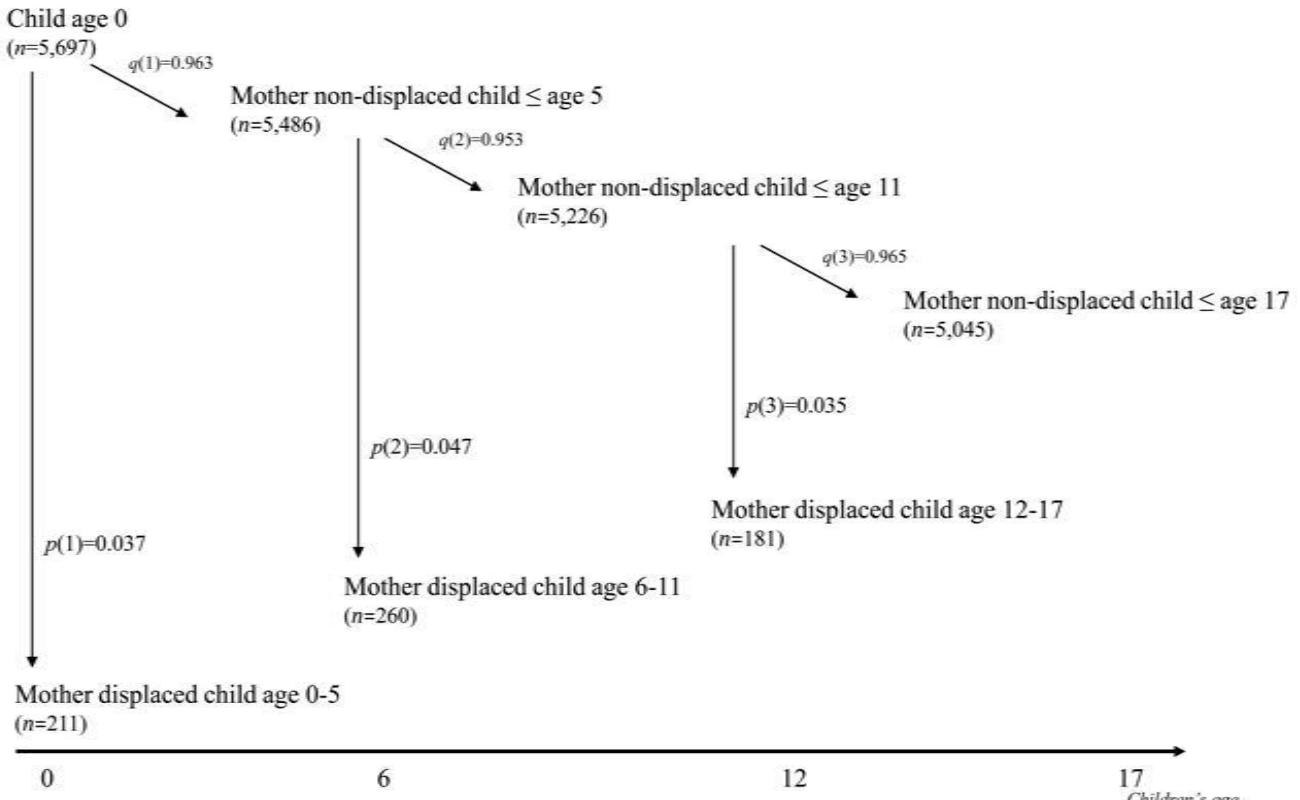

FIG. 1.—Time-varying treatment flow chart: maternal job displacement by child age

RESULTS

Descriptive Statistics

Table 1 describes covariates used to predict maternal displacement and children's outcomes.[15] Displaced mothers are more disadvantaged than nondisplaced mothers on most indicators: they have more disadvantaged family backgrounds, have lower measured ability and fewer years of education, are more likely to have engaged in delinquent activity, have more children, have less employer tenure, and are more likely to be in manufacturing and trade industries. We also report descriptive statistics for our five outcomes. Children whose mothers were displaced have lower educational attainment and higher levels of depressive symptoms than children whose mothers were not displaced.

Effects of Maternal Displacement on Children's Outcomes

We first derived estimated propensity scores for maternal displacement for children ages 0–17 using the set of predisplacement covariates. Logistic regression results reported in table 2 suggest that black mothers are significantly more likely to be displaced than white mothers, as are mothers with low measured ability, mothers who have engaged in delinquent activity, and mothers who do not attend college. Employer tenure is protective as is working in a nonmanufacturing and nontrade industry. Our propensity equation does not have strong predictive ability, as expected given the relatively exogenous shock of a displacement event (Brand 2004).[16] Table 2 reports only one of the nine propensity score models we run, given that we also estimate multistate and time-varying displacement effects. The alternative propensity model estimates are available on request.

We report unmatched differences and two sets of propensity score matching estimates of maternal displacement on children's outcomes in young adulthood in table 3. Results pertain to maternal displacement that occurs anytime when the child is ages 0–17.[17] The unmatched differences

[15] We impute missing values using the full set of covariates for all variables with at least 1% missing. Mothers' mothers' education (i.e., children's grandmothers), mothers' delinquency, and mothers' ability had the most missing values (7.6%, 6.6%, and 4.8%, respectively). The remaining missing propensity score values (2.8%, $n = 158$) were imputed using the results of successive propensity score models for each year of the survey.

[16] Alternatively, we are missing covariates that would enable stronger prediction of maternal displacement events.

[17] In results not shown, we stratified analyses by a child's gender and found that effect estimates do not significantly differ. We also ran regression models in which we cluster on mothers to account for the dependence among siblings; substantive findings were not affected by this adjustment.

TABLE 2

LOGIT REGRESSION ESTIMATES OF MATERNAL JOB DISPLACEMENT DURING CHILDHOOD ON PREDISPLACEMENT COVARIATES

Maternal time-invariant covariates:	
Black (0/1)	.441***
	(.110)
Hispanic (0/1)	.235
	(.144)
(Grand)mother's years of education	−.002
	(.017)
(Grand)parents' intact family age 14 (0/1)	−.005
	(.089)
(Grand)parents' southern residence (0/1)	−.023
	(.118)
Mental ability (ASVAB; −3 to 3)	−.384***
	(.077)
Rotter locus of control (6–14)	.017
	(.019)
Delinquency (0/1)	.308**
	(.106)
Maternal covariates, child age 0:	
Age	−.045*
	(.022)
Number of children	−.031
	(.040)
Marital or cohabiting partner (0/1)	−.481***
	(.093)
High school completion (0/1)	.371**
	(.111)
College completion (0/1)	−2.728**
	(1.010)
Region (0/1)	−.011
	(.112)
Employer tenure (weeks)	−.005*
	(.002)
Full-time employment (0/1)	.723
	(.743)
Manufacturing worker (0/1)	.480***
	(.138)
Trade worker (0/1)	.044
	(.123)
Child covariate:	
Child age in 2010	−.041*
	(.020)
Constant	−1.310
	(1.273)
Logit regression χ^2	166.46
$P > \chi^2$	.000
n	5,697

NOTE.—Numbers in parentheses are SEs. Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and who were at least 19 years old in 2010. Two-tailed tests.

$^{+}P < .10$.

* $P < .05$.

** $P < .01$.

*** $P < .001$.

TABLE 3
MATCHING ESTIMATES OF MATERNAL JOB DISPLACEMENT DURING CHILDHOOD ON CHILDREN'S OUTCOMES IN YOUNG ADULTHOOD

	Unmatched Differences	Propensity Score Kernel Matching, TT	Propensity Score Kernel Matching, TUT
Educational attainment:			
High school completion (0/1) . . .	−.053**	−.037^{+}	−.052*
	(.021)	(.022)	(.024)
College attendance (0/1)	−.089***	−.063**	−.099***
	(.023)	(.023)	(.022)
College completion (0/1)	−.051**	−.036*	−.043*
	(.019)	(.016)	(.017)
Psychological well-being:			
CESD ages 20–24 (0–1)	.018^{+}	.014	.022*
	(.009)	(.009)	(.010)
CESD ages 25–29 (0–1)	.030**	.025*	.034**
	(.011)	(.011)	(.012)

NOTE.—Numbers in parentheses are SEs. Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and who were at least 19 years old in 2010. Maternal displacement is measured as displacement that occurred when children were 0–17 years old. Propensity scores were estimated by a logit regression model of maternal displacement on the set of predisplacement covariates. TUT SEs were bootstrapped on the basis of 50 replications. Two-tailed tests.

$^{+}P < .10$.

* $P < .05$.

** $P < .01$.

*** $P < .001$.

establish a benchmark to compare to matched results. All effect estimates are statistically significant. Children whose mothers were displaced have a 5 percentage point lower level of high school and of college completion (or a 22% and 41% decrease in the expected odds, respectively) and a 9 percentage point lower level of college attendance (or a 32% decrease in the expected odds) than children whose mothers were not displaced. Our measure of CESD ranges from 0 to 1, with an overall mean of about 0.23 for both 20–24- and 25–29-year-olds. Children whose mothers were displaced have a 0.02 increase in depressive symptoms at ages 20–24 and a 0.03 increase at ages 25–29, relative to children whose mothers were not displaced.

We next turn to kernel-matching results.[18] We report both treatment effects for the treated (TT; i.e., effects pertaining to children whose moth-

[18] We also estimate t-tests for equality of means between treated and control groups, the standardized bias before and after matching, and the achieved percentage reduction in bias. The standardized bias is the difference in means in the treated and the nontreated (full or matched) subsamples as a percentage of the square root of the average of the

ers were displaced) and treatment effects for the untreated (TUT; i.e., effects pertaining to children whose mothers were not displaced, had they been displaced) to provide preliminary evidence as to variation in effects by the propensity for treatment. For the TT, although point estimates of maternal displacement effects are lower than unmatched differences, displacement remains associated with lower levels of high school completion and college attendance and completion (a 4, 6, and 4 percentage point difference, or a 15%, 24%, and 33% decrease in the expected odds, respectively) and higher levels of depressive symptoms in the late 20s. The estimates of the TUT are larger than those for the TT for all five outcomes, including a 22%, 35%, and 35% decrease in the odds of high school completion, college attendance, and college completion, respectively, and larger increases in depressive symptoms. As the distribution of the untreated population is weighted toward lower-propensity individuals, TUT > TT suggests that effects of maternal displacement on children's educational attainment and psychological well-being in young adulthood are larger for children whose mothers are unlikely to experience a displacement event.

Variation in Effects of Maternal Displacement on Children's Outcomes

Heterogeneous maternal displacement effects.—Average effects of maternal displacement on children's outcomes may conceal underlying systematic effect heterogeneity (i.e., variation in effects by selection into treatment) shaped by the population composition of children of displaced mothers (Xie et al. 2012; Brand and Simon Thomas 2013). To assess effect heterogeneity we use the SM method, in which we generate balanced propensity score strata, estimate effects by strata (level 1), and then estimate the trends in effects (level 2). In contrast to comparing differences between TT and TUT matching estimates, we explicitly estimate effects across the propensity score distribution and assess the trend in effects using the SM approach. Our analysis resulted in five strata by which the sample is divided, where stratum 1 corresponds to the lowest propensity and stratum 5 to the highest propensity children (on the basis of their mothers' probability of dis-

sample variances in the treated and the control groups (Rosenbaum and Rubin 1985). We achieve substantial reduction in bias via matching (results available on request). Morgan and Winship (2007) find that kernel matching (Leuven and Sianesi 2003) performs best (has least bias) with a well-specified propensity score equation but that matching with five nearest neighbors (Abadie et al. 2004) performs best with a poorly specified propensity score equation. In results not shown, we compare kernel-matching estimates to matching with five nearest neighbors as a sensitivity test for misspecification. Results are substantively similar.

placement).[19] As we observe from the stratum-specific population composition of children and their displaced mothers reported in appendix B, young black mothers with low measured ability who are high school dropouts, engaged in delinquent activity, grew up in the south, work in manufacturing, and have only a few months job tenure are characteristic of mothers with a high propensity for displacement. By contrast, relatively older white mothers with higher measured ability who attend college, work outside manufacturing, and have at least a year of job tenure are characteristic of mothers with a low propensity for displacement.

Results reported in table 4 suggest declines in the deleterious effects of maternal displacement as the propensity for displacement increases. That is, examining the trend in treatment effects (i.e., the level 2 slopes), we find that children with the most advantaged mothers who have the lowest propensity for displacement have the largest negative effects of maternal displacement on high school completion and college attendance and the largest positive effects on depressive symptoms outcomes in young adulthood. We find, however, no trend in effects for college completion. We observe several significant level 1 slopes among low-propensity children in stratum 1—generally larger than the average effects we report in table 3—one marginally significant effect in stratum 2, and no significant effects among higher-propensity children in strata 3–5. Figures 2*A* and 2*B* summarize the results from table 4. The *x*-axes indicate propensity score strata, and *y*-axes the displacement effects, where plotted symbols represent level 1 and lines represent level 2 slopes. Results from table 4 and figure 2 provide further support for the matching results in table 3 suggesting TUT > TT, as both indicate that children with a low propensity for maternal displacement are most negatively affected by the event.

As we do not observe a clear monotonic trend in effects, we test the sensitivity of our analyses to the parametric assumption we impose in SM using the MS heterogeneous treatment effects method. Figures 3*A* and 3*B* depict local polynomial smoothed curves fit to kernel-matched differences between children whose mothers were and were not displaced by estimated propensity scores. The *x*-axes indicate estimated propensity for maternal displacement, and *y*-axes the matched differences between treated and controls. The scatter plots are omitted for simplicity. Confirming our results from SM, we observe that the negative effect of maternal displacement on high school completion and college attendance and the positive effect on depressive symptoms decrease as the propensity for displacement increases. The effects on college completion have an approximate U-shape,

[19] Even the highest propensity score stratum has a lower-bound propensity score of 0.2 on a range of 0–1, due to the limited predictive ability of our propensity score equation.

TABLE 4

MATCHING ESTIMATES OF HETEROGENEOUS MATERNAL JOB DISPLACEMENT ON CHILDREN'S OUTCOMES IN YOUNG ADULTHOOD: STRATIFICATION MULTILEVEL METHOD (SM-HTE)

	LEVEL 1					LEVEL 2
	Stratum 1 $p = [0–.1)$	Stratum 2 $p = [.1–.15)$	Stratum 3 $p = [.15–.175)$	Stratum 4 $p = [.175–.2)$	Stratum 5 $p = [.2–1)$	Trend
Educational attainment:						
High school completion (0/1)	−.101**	.007	−.094	−.095	.080	.024
	(.038)	(.038)	(.061)	(.060)	(.058)	(.015)
College attendance (0/1)	−.132**	−.073^{+}	−.022	−.067	.024	.033^{+}
	(.045)	(.040)	(.068)	(.066)	(.063)	(.017)
College completion (0/1)	−.051	−.023	.022	−.071	−.038	−.002
	(.041)	(.029)	(.054)	(.049)	(.039)	(.012)
Psychological well-being:						
CESD ages 20–24 (0–1)	.043*	−.003	.027	−.001	.005	−.008
	(.017)	(.016)	(.025)	(.025)	(.025)	(.007)
CESD ages 25–29 (0–1)	.051**	.024	.029	.012	−.013	−.014^{+}
	(.019)	(.018)	(.028)	(.030)	(.028)	(.007)
n	2,541	1,778	597	379	402	5,697

NOTE.—Numbers in parentheses are SEs. Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and who were at least 19 years old in 2010. Maternal displacement is measured as displacement that occurred when children were 0–17 years old. Propensity scores were estimated by a logit regression model of maternal displacement on the set of predisplacement covariates. Two-tailed tests.

$^{+}P < .10$.

* $P < .05$.

** $P < .01$.

*** $P < .001$.

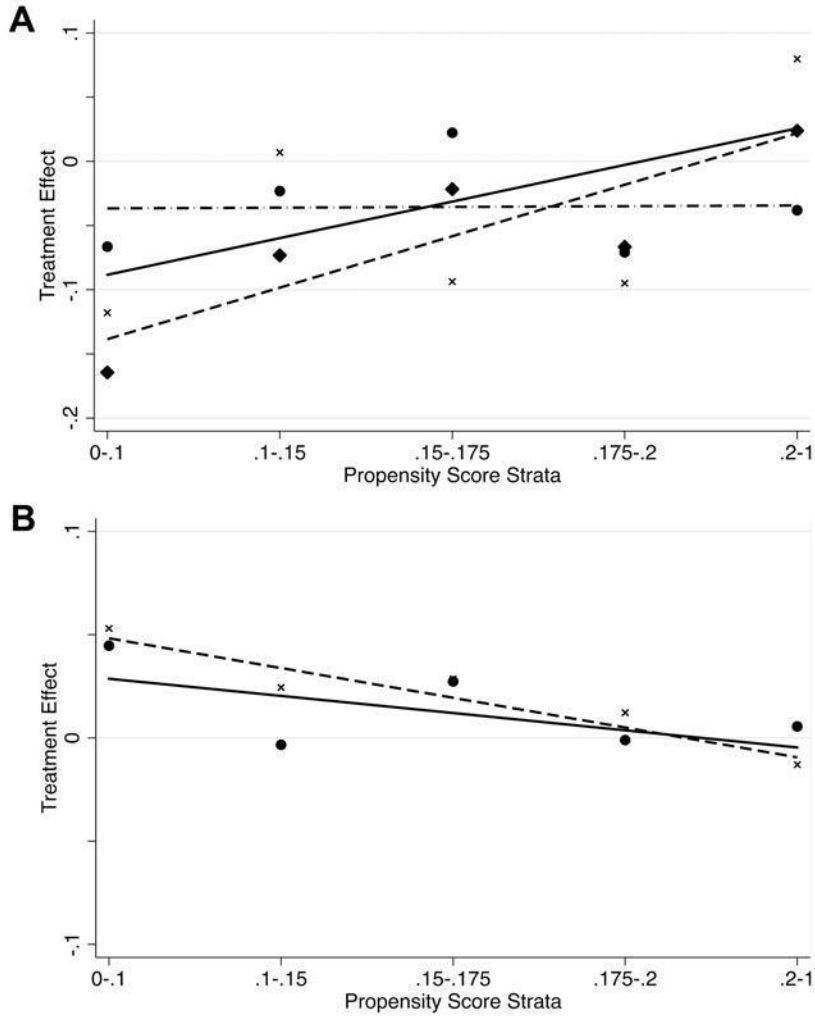

FIG. 2.—Stratification-multilevel heterogeneous maternal displacement on children's educational attainment (*A*) and social-psychological well-being (*B*): *crosses* and *solid lines*, high school completion; *diamonds* and *dashed lines*, college attendance; *circles* and *dot-dashed lines*, college completion.

indicating the largest effect for children whose mothers have a midpropensity for displacement.[20]

[20] To facilitate implementation of our analyses of heterogeneous treatment effects, we use the Stata modules hte for SM and hte2 for MS (Jann et al. 2010*a*, 2010*b*). We thank Ben Jann for programming modifications to hte2 to facilitate presentation of our results.

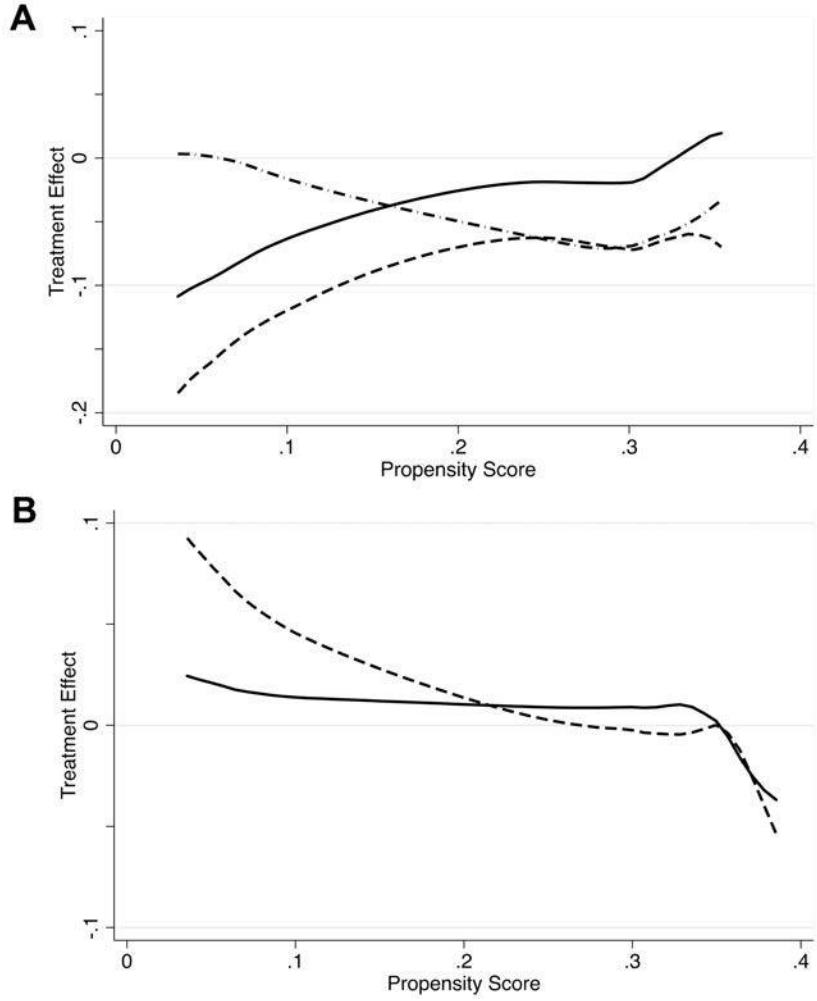

FIG. 3.—Matching-smoothing heterogeneous maternal displacement on children's educational attainment (*A*) and social-psychological well-being (*B*): *solid lines*, high school completion; *dashed lines*, college attendance; *dot-dashed lines*, college completion.

We do not mean to suggest that children with mothers who have a high propensity of displacement are more advantaged in their levels of educational attainment and psychological well-being, as they nevertheless face a variety of disadvantaged circumstances through childhood; they are, nonetheless, less affected by maternal displacement compared to lower-propensity children. To underscore this point, we report levels of children's outcomes by maternal displacement and strata in appendix B.

Multistate maternal displacement effects.—In table 5, we report estimates of multistate maternal displacement effects on children's outcomes in young adulthood. We compare effects for children whose mothers were displaced during recessionary and nonrecessionary periods (to assess differences in macroeconomic contexts) and effects for children whose mothers were displaced in LMAs with at least 9% unemployment to areas with less than 9% unemployment (to assess differences in local contexts). We also compare the TT to the TUT within each treatment condition. For the TT, we find a significant 8 percentage point lower level of high school completion and a 6 percentage point lower level of college completion (or a 29% and 49% decrease in the expected odds, respectively) for children whose mothers were displaced during recessions. The TUT estimate is larger for high school completion than the TT estimate and significant for college attendance but smaller for college completion. We find no significant effects on depressive symptoms for children whose mothers were displaced during recessions. For children whose mothers were displaced in nonrecessionary periods, with one exception for the TT estimate for high school completion, effects estimates are significant for all outcomes. Point estimates are somewhat smaller for high school and college completion but larger for college attendance, suggesting a 7 and 11 percentage point lower level of college attendance for the TT and the TUT, respectively. We find significant effects on depressive symptoms throughout the 20s, with larger effects in the late than the early 20s and larger effects for the TUT than the TT.

Results for children whose mothers were displaced in LMAs with high and nonhigh unemployment are similar to those for recessionary and nonrecessionary displacements. That is, only high school completion and college attendance (for the TUT) are significant among children whose mothers were displaced in high-unemployment areas, whereas all outcomes (except high school completion for the TT) are significant for children whose mothers were displaced in areas without high levels of unemployment. Point estimates are likewise similar across the two sets of results. Yet it is not the case that these are the same children across these treatment conditions (see app. C). Mothers displaced in high-unemployment areas are more disadvantaged relative to mothers displaced in recessionary periods. Moreover, low-propensity mothers differ from mothers displaced in low-unemployment contexts (where in both cases, we find the largest effects of displacement): family backgrounds and employment profiles substantially differ, with low-propensity mothers more advantaged. These groups share in common, however, that the displacement event should be a greater shock and potentially incite a greater sense of deprivation relative to high-propensity displaced mothers and mothers displaced in high-unemployment contexts, where economic adversity deviates less from the social norm.

TABLE 5

MATCHING ESTIMATES OF MULTISTATE MATERNAL JOB DISPLACEMENT ON CHILDREN'S OUTCOMES IN YOUNG ADULTHOOD

	MOTHER DISPLACED							
	RECESSIONARY PERIOD		NONRECESSIONARY PERIOD		LMA WITH HIGH UNEMPLOYMENT		LMA WITHOUT HIGH UNEMPLOYMENT	
	TT	TUT	TT	TUT	TT	TUT	TT	TUT
Educational attainment:								
High school completion (0/1)	−.075*	−.083^{+}	−.037	−.050^{+}	−.092^{+}	−.098^{+}	−.024	−.044^{+}
	(.037)	(.042)	(.024)	(.028)	(.043)	(.052)	(.024)	(.023)
College attendance (0/1)	−.056	−.074*	−.070**	−.105***	−.051	−.079^{+}	−.073**	−.105**
	(.039)	(.037)	(.025)	(.023)	(.045)	(.043)	(.025)	(.034)
College completion (0/1)	−.061*	−.050	−.035*	−.044*	−.008	−.020	−.049**	−.056**
	(.026)	(.035)	(.018)	(.017)	(.035)	(.046)	(.017)	(.019)
Psychological well-being:								
CESD ages 20–24 (0–1)	.005	.016	.019^{+}	.022*	.002	.006	.020^{+}	.028*
	(.014)	(.015)	(.010)	(.011)	(.017)	(.019)	(.010)	(.012)
CESD ages 25–29 (0–1)	.000	.010	.035**	.049***	.006	.024	.031*	.041**
	(.016)	(.015)	(.012)	(.013)	(.018)	(.021)	(.012)	(.014)

NOTE.—Numbers in parentheses are SEs. Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and children who were at least 19 years old in 2010. Propensity scores were estimated by a series of logit regression models of the specific state of maternal displacement on the set of predisplacement covariates. Recessionary periods include 1990–92, 2001–2, and 2008–10; nonrecessionary periods include the remaining periods. High-unemployment areas include any displacement that occurred when the respondent lived in a labor market area (LMA) with at least 9% unemployment; without-high-unemployment areas include those with less than 9% unemployment. TUT SEs were bootstrapped on the basis of 50 replications. Two-tailed tests.

$^{+}P < .10$.

* $P < .05$.

** $P < .01$.

*** $P < .001$.

Time-varying maternal displacement effects.—We report TT and TUT kernel-matching estimates of time-varying effects of maternal displacement in table 6. For columns labeled "mother displaced child ages 0–5," we matched by propensity scores estimated with a logit regression of maternal displacement when children are 0–5 on the time-invariant predictors and those predictors corresponding to children at age 0. In columns labeled "mother displaced child ages 6–11," we used the same time-invariant predictors and predictors corresponding to children at age 6 to construct propensity scores. And for "mother displaced child ages 12–17," we used the time-invariant predictors and predictors corresponding to children at age 12. As we describe above, children whose mothers were displaced in prior periods are no longer at risk for initial maternal displacement and were dropped from subsequent comparisons.[21]

We find only one marginally significant estimate of maternal displacement occurring in early childhood (TUT for college attendance).[22] By contrast, we find significant deleterious effects of maternal displacement occurring in middle childhood on college attendance (a 30% decrease in the expected odds) and psychological well-being in the late 20s (in both cases the TUT is modestly larger than the TT). For children whose mothers were displaced when they were adolescents, we find large, significant TT effects on every outcome. The TUT point estimates are all greater than the TT for adolescents whose mothers were displaced, but effects on college completion and depressive symptoms in the late 20s are not significant. For the TT, we observe a 12 percentage point lower level of high school completion, a 10 percentage point lower level of college attendance, and a 5 percentage point lower level of college completion (a 40%, 25%, and 45% decrease in the expected odds, respectively). We find large effects of displacement on depressive symptoms in the 20s among children whose moth-

[21] Our time-varying treatments were constructed such that adolescents whose mothers experienced displacement were preceded by a decade or more in which no job loss occurred, while young children whose mothers experienced displacement may have been preceded by frequent losses. In results not shown, we estimated two additional series of matching models to test the sensitivity of our results to this specification. First, we estimated time-varying maternal displacement effects in which we eliminated children whose mothers experienced multiple displacements (47 cases for children ages 0–5, 63 casesfor children ages 6–11, and 15 cases for children ages 12–17). Second, we estimated time-varying effects in which we relaxed our preferred event history setup and allowed the control units to include children whose mothers were displaced at earlier periods (resulting in no change to the sample of children ages 0–5 but adding 23 cases to children ages 6–11 and 53 cases to children ages 12–17). Both specifications yield results that are substantively similar to our original time-varying specification.

[22] Young children have a longer period between the displacement event and the outcome than do older children whose mothers were displaced. It is unclear, however, whether maternal displacement effects would diminish or accumulate over this period.

TABLE 6
MATCHING ESTIMATES OF TIME-VARYING MATERNAL JOB DISPLACEMENT ON CHILDREN'S OUTCOMES IN YOUNG ADULTHOOD

	MOTHER DISPLACED					
	CHILD 0–5		CHILD 6–11		CHILD 12–17	
	TT	TUT	TT	TUT	TT	TUT
Educational attainment:						
High school completion (0/1)	−.004	−.006	−.016	−.031	−.115**	−.137***
	(.037)	(.044)	(.032)	(.033)	(.040)	(.037)
College attendance (0/1)	−.032	−.070^{+}	−.082*	−.085*	−.099*	−.108**
	(.042)	(.036)	(.034)	(.036)	(.039)	(.041)
College completion (0/1)	−.035	−.031	−.027	−.023	−.052^{+}	−.057
	(.031)	(.032)	(.024)	(.027)	(.027)	(.037)
Psychological well-being:						
CESD ages 20–24 (0–1)	.003	.009	.013	.013	.031*	.040*
	(.017)	(.019)	(.013)	(.013)	(.015)	(.016)
CESD ages 25–29 (0–1)	−.021	−.027	.047**	.056**	.034^{+}	.037
	(.016)	(.017)	(.017)	(.021)	(.019)	(.024)

NOTE.—Numbers in parentheses are SEs. Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and who were at least 19 years old in 2010. Propensity scores were estimated by a series of logit regression models of maternal displacement on the set of predisplacement time-invariant and time-varying covariates. TUT SEs were bootstrapped on the basis of 50 replications. Two-tailed tests.

$^{+}P < .10$.
* $P < .05$.
** $P < .01$.
*** $P < .001$.

ers were displaced in adolescence, but effects in the late 20s are larger among children whose mothers were displaced in middle childhood.

Descriptive statistics of family and socioeconomic statuses of the displaced.—Table 7 reports descriptive statistics of family and socioeconomic statuses pre- and postmaternal displacement and throughout childhood, for our full sample of children whose mothers were displaced and each subpopulation defined by displacement propensity, context, and timing. Panel A describes family structure heterogeneity over the course of childhood for the population of children whose "single" mothers were displaced. The average child of a displaced mother spends 18% of childhood with married parents. Children of displaced mothers have, on average, two maternal relationship transitions (between marriage or partnership states). We observe little difference in marital or partnership statuses, or the number of

TABLE 7
DESCRIPTIVE STATISTICS OF FAMILY AND SOCIOECONOMIC STATUSES OF THE DISPLACED

	Overall	Displacement Propensity					Displacement Context				Displacement Timing		
	Child 0–17	p = [0–.1)	p = [.1–.15)	p = [.15–.175)	p = [.175–.2)	p = [.2–1)	Recession	Nonrecessionary	High Unemployment	Without High Unemployment	Child 0–5	Child 6–11	Child 12–17
A. Family status:													
Married:													
Postdisplacement (0–2 years)	.129	.228	.153	.066	.047	.032	.083	.146	.177	.113	.237	.077	.070
Postdisplacement (2–4 years)	.166	.221	.191	.132	.141	.064	.124	.171	.221	.147	.251	.146	.083
Proportion throughout childhood	.184	.298	.205	.108	.091	.072	.184	.181	.213	.174	.217	.168	.168
Partnered:													
Predisplacement (0–2 years)	.169	.304	.186	.118	.082	.021	.124	.181	.114	.188	.161	.173	.173
Postdisplacement (0–2 years)	.182	.316	.181	.105	.129	.064	.161	.175	.152	.192	.137	.211	.192
Postdisplacement (2–4 years)	.134	.202	.135	.079	.106	.086	.150	.132	.095	.147	.109	.138	.160
Proportion throughout childhood	.103	.161	.104	.086	.072	.036	.088	.105	.091	.107	.092	.114	.099
Number of relationship transitions	2.109	2.634	2.242	1.848	1.671	1.441	2.128	2.057	1.757	2.085	2.104	2.192	1.994

B. Socioeconomic status:													
Unemployment:													
Postdisplacement (0–2 years)	.844	.854	.837	.868	.819	.848	.880	.832	.880	.832	.813	.842	.890
Postdisplacement (2–4 years)	.495	.355	.598	.547	.374	.565	.479	.497	.456	.508	.525	.494	.454
Proportion throughout childhood	.408	.338	.407	.451	.425	.484	.395	.419	.397	.411	.342	.444	.432
Welfare:													
Predisplacement (0–2 years)	.506	.331	.493	.676	.712	.516	.432	.546	.538	.494	.500	.537	.460
Postdisplacement (0–2 years)	.555	.446	.546	.640	.658	.604	.495	.582	.590	.543	.512	.607	.527
Postdisplacement (2–4 years)	.482	.374	.483	.575	.500	.571	.449	.503	.490	.479	.458	.524	.443
Proportion throughout childhood	.368	.272	.347	.442	.442	.465	.337	.387	.364	.369	.292	.404	.404
Wages:													
Predisplacement (0–2 years)	6.302	8.841	6.182	4.961	4.839	5.028	8.053	5.591	6.051	6.389	5.058	5.409	9.472
Postdisplacement (0–2 years)	2.095	1.840	2.420	1.724	2.197	1.990	1.889	2.184	1.610	2.260	2.313	2.033	1.907
Postdisplacement (2–4 years)	5.779	7.118	4.916	5.507	6.470	5.089	6.390	5.737	6.775	5.434	5.384	5.542	6.746

NOTE.—Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and who were at least 19 years old in 2010 ($n = 5{,}697$).

relationship transitions, among mothers across displacement contexts. We find higher postdisplacement marriage levels, but not partnership levels or relationship transitions, among mothers who were displaced when children were young, relative to mothers who were displaced when children were older. The most notable differences lie across displacement propensity. Children whose mothers have a higher likelihood of displacement spend less time with married or partnered parents and have more relationship transitions. Low-propensity children are therefore more likely to have another resident adult present but nevertheless have larger effects of maternal displacement on attainment and well-being.

In table 7 panel B, we describe unemployment, wages (among employed workers), and "welfare" receipt (including welfare, food stamps, or other public assistance). Postdisplacement unemployment is uniformly high across displaced subpopulations, but particularly for those displaced in high-unemployment contexts. Wages significantly decline after a displacement but generally resume predisplacement levels two to four years post. Mothers displaced in recessions have larger wage declines than mothers displaced in expansions, but we find little difference between mothers displaced in high relative to nonhigh unemployment areas. Mothers displaced when children are adolescents have higher predisplacement wages (as these mothers tend to be older) and have greater wage declines postdisplacement than mothers displaced when children are younger. As expected, mothers' predisplacement wages decrease and levels of welfare increase as the propensity for displacement increases. Wages in the period immediately after displacement are roughly equalized across the propensity for displacement but return to their predisplacement pattern two to four years post. We thus observe a potentially larger short-term economic impact, but less overall economic adversity, among low-propensity mothers.

CONCLUSION

The study of parental job displacement effects on children augments an extensive literature on the relationship between family socioeconomic status and children's life outcomes, by relating career shocks, economic restructuring, and structural labor market conditions to parental attainment and the intergenerational transmission of status. Job displacement is a precipitating life event that entails a sequence of stressful experiences, from job loss notification, anticipation, dismissal, and unemployment to job search, possible retraining, and eventual reemployment, often in a job of inferior quality and lower earnings relative to the job lost. Displacement is the result of conditions that are largely beyond the control of parents, nonetheless hindering parents' ability to sustain a career characterized by upward mobility and to transmit accrued advantages to children. By studying displace-

ment and its effects on intergenerational mobility, we capitalize on a scientific opportunity provided by extreme economic change. Abrupt changes in economic conditions provide a stronger basis for the study of socioeconomic transmission than the usual practice of examining the covariation of outcomes with socioeconomic status that arise from a variety of sources over an indeterminate period.

Despite this motivation for studying displacement and the large literature linking displacement to deleterious worker outcomes, we know relatively less about how displacement among parents, particularly single parents, affects children. The possibility that children will endure adverse socioeconomic consequences and model psychological distress resulting from displacement is heightened in single-parent families because such children generally lack the advantage of an additional parent or parental figure who may either be employed or possibly temper a depressive outlook with a more optimistic and positive style of coping. As past research on displacement is largely silent as to intergenerational effects among single parents, so too past research on single parents is fairly limited as to the consequences of involuntary employment separations. Much of this latter work treats employment status as a static state, and those studies attending to employment trajectories among single parents treat some combination of voluntary and involuntary transitions as employment instability, thereby limiting causal inference.

In this study, we expand and integrate past research by examining how job displacement among single mothers divides children's educational and social-psychological outcomes in young adulthood. We use data on women from the NLSY who were followed for the last three decades and on their children from the NLSCM who have been followed for over two decades. The NLSY and NLSCM are well suited for studying intergenerational effects of maternal displacement. They contain rich panel data on the social background, education, and labor market experiences of women and their children. In contrast to most panel data, the NLSY does not group workers who lost jobs because of layoffs with workers who were fired; consequently, selection bias poses less of an obstacle to estimating effects of involuntary employment separation. Estimating the propensity for maternal displacement with a range of maternal characteristics, we find that mothers with disadvantaged socioeconomic backgrounds, with low achievement and ability, who engaged in delinquent activity, and with low adult socioeconomic status are most likely to be displaced from jobs and consequently face an additional hardship while raising their children. Using propensity score kernel matching, we find significant deleterious effects of maternal displacement on children's high school completion, college attendance and completion, and depressive symptoms as young adults. These lower levels of educational attainment and social-psychological well-being render children of

displaced mothers more likely to experience job displacement over the course of their own careers, signifying a potential intergenerational transmission of employment instability.

We also explore variation in the effects of displacement by the likelihood a child's mother lost her job, by loss occurring in contexts with more or less unemployment, and by the child's age at which a mother lost her job (i.e., by displacement propensity, context, and timing, respectively). As to variation in effects by displacement propensity and context, one possibility is that income shocks are more damaging to children of families with a high likelihood of displacement and those who experience the event in high-unemployment contexts, that is, situations in which economic adversity tends to be greater (von Wachter 2010; Couch et al. 2011; Davis and von Wachter 2011). But this conjecture is largely unsupported by our analyses of the effects of maternal displacement on children, particularly for effects on children's social-psychological well-being. Our results instead suggest that maternal displacement among children whose mothers had a low likelihood of displacement and children whose mothers were displaced in low-unemployment contexts yield the largest effects. Prior research found larger effects on workers' social-psychological distress in contexts without widespread unemployment, as such settings heighten the internalization of blame and social stigma associated with job loss (Miller and Hoppe 1994; Clark 2003, 2010; Charles and Stephens 2004; Brand et al. 2008), but these studies have not considered effects on children. Yet another strand of research demonstrates that social stigma and relative deprivation are primary factors linking socioeconomic shocks to child well-being (MacLeod and Shanahan 1993; Duncan and Brooks-Gunn 1997). Thus, although economic adversity is generally less for children of low-propensity mothers and those displaced in lower-unemployment contexts, such families lack referents to similarly strained families and a social norm of deprivation. Mothers who have a high likelihood of displacement, by contrast, may expect more socioeconomic instability in their lives and be embedded in a social network in which income shocks and economic distress are normative experiences and less stigmatized as a result of individual failure, rendering the effects of displacement on their children less severe.

As to variation in displacement effects by timing, despite developmental theory suggesting critical consequences of socioeconomic adversity in early childhood, we find no negative effects among young children whose mothers were displaced. We find, however, significant effects when maternal displacement occurs in middle childhood and, especially, adolescence. To the extent that older children are more sensitive to experiences of social stigma and relative deprivation than younger children, these findings may likewise suggest the importance of these social-psychological mechanisms

in the impact of family socioeconomic shocks on child well-being. There are a few alternative, or additional, explanations for variation in effects by displacement timing. The positive effects of mothers who are home to tend to their young children may counterbalance the negative effects and economic pressures associated with job loss. Or the lengthy time elapsing between maternal displacement in early childhood and young adult outcomes may dilute, rather than strengthen, initial effects (at least for educational attainment). Or adolescents may enter the labor market when mothers lose jobs to partially offset family economic distress, and thus be less likely to continue their education.

We caution that results such as ours are always subject to the possibility that some important omitted variables differentiate children whose mothers were and were not displaced. Indeed, an alternative interpretation for heterogeneity in effects involves differential selection bias. That is, women with a low propensity for job displacement and women displaced in relatively low-unemployment settings may have more unobservable characteristics that bias the relationship between maternal displacement and children's outcomes. We cannot adjudicate between this possibility and the explanations we offer above, but we speculate that both play a role. Our study is also limited in that we do little to attend to the heterogeneity of the "single mother" population. Future work will explore the complex interrelationship among family structure, parental displacement, and child well-being. Moreover, we do not distinguish different recessionary periods. As data become available, future work should focus more explicitly on the effects of parental displacement occurring during the Great Recession on children's long-term outcomes. Finally, we do not explicitly assess potential mechanisms linking maternal displacement to children's outcomes. Our approach focuses on first determining that main effects exist, an issue that, as demonstrated throughout these analyses, is sufficiently complicated and important as to warrant undivided attention. The effects of mediators on later outcomes, often estimated by simply including such variables in expanded regression models, seldom warrant causal interpretations and can lead to erroneous conclusions regarding both the intermediary and main effects (Holland 1988; Sobel 2008; Elwert and Winship 2013). Future work should attend to the causal effects of the mechanisms linking parental displacement to children's well-being.

As at least half of all children will spend some portion of their childhood raised by a single mother, the socioeconomic well-being of such families is a fundamental concern. We should protect disadvantaged children because they have not made the choices that have resulted in their socioeconomic conditions, or so goes the rhetoric on social class disparities in children's resources. Such discourse implicitly assumes mothers have made

such choices. But women are also subject to structural conditions largely beyond their control. Debates about social assistance should acknowledge that job separation among single mothers is at times involuntary and that such involuntary events are associated with long-term unemployment, socioeconomic and social-psychological decline, and significant intergenerational effects. We should restrict assistance neither to the most disadvantaged mothers nor to those mothers only displaced in economic contractions, as particularly deleterious maternal displacement effects on life trajectories of children may accrue among otherwise more advantaged single-parent families.

APPENDIX A

All matching estimators of the TT take the following general form:

$$\mathrm{TT} = \frac{1}{n_1} \sum_{i}^{n_i} \left(y_{i,d=1} - \sum_{i(j)}^{ij} w_{i(j)} y_{i(j),d=0} \right), \tag{A1}$$

where n_1 is the number of treatment cases, i is the index over treatment cases, $i(j)$ is the index over untreated cases for treated case i ($i(j) = 1, \ldots i(J)$), and $w_{i(j)}$ is the scaled weight (with sum of 1) that measures the relative importance of each untreated case. The estimator for the TUT is simply the corollary, where each untreated case is matched to a treated case. Using kernel matching, all untreated units are used and weighted according to the distance from the estimated propensity score of the target treated unit. We define the kernel-matching weight as

$$w_{i(j)} = \frac{G\left[\dfrac{P(s_j) - P(s_i)}{a_n}\right]}{\sum_j G\left[\dfrac{P(s_j) - P(s_i)}{a_n}\right]}, \tag{A2}$$

where a_n is a bandwidth parameter that scales the difference in the estimated propensity scores on the basis of the sample size and $\hat{P}(\cdot)$ is the estimated propensity score. While scholars have not reached a consensus as to which matching estimator performs best in each application or which expected variance of matching estimates should be used (Morgan and Harding 2006; Morgan and Winship 2007), Morgan and Winship (2007) find that kernel matching (Leuven and Sianesi 2003) has the least bias with a well-specified propensity equation.

APPENDIX B

TABLE B1
COVARIATE MEANS BY MATERNAL DISPLACEMENT AND PROPENSITY SCORE STRATA

	Stratum 1 $p = [0–.05)$			Stratum 2 $p = [.05–.10)$			Stratum 3 $p = [.10–.15)$			Stratum 4 $p = [.15–.20)$			Stratum 5 $p = [.20–1)$		
	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B
Maternal time-invariant covariates:															
Black (0/1)	.080	.101	.016	.316	.278	.012	.467	.564	.010	.854	.928	.120	.770	.967	.103
Hispanic (0/1)	.107	.053	.053	.124	.079	.019	.059	.119	.075	.049	.072	.155	.018	.033	.130
(Grand)mother's years of education	10.987	11.610	.102	9.696	9.860	.060	10.084	10.037	.239	10.262	10.125	.094	9.504	10.081	.161
(Grand)parents' intact family age 14 (0/1)	.697	.720	.079	.611	.630	.052	.539	.504	.004	.579	.695	.227	.503	.521	.048
(Grand)parents' southern residence (0/1)	.297	.178	.002	.486	.502	.103	.362	.343	.015	.612	.639	.179	.623	.658	.272
Mental ability (ASVAB; −3 to 3)	.026	.004	.257	−.429	−.764	.017	−.642	−.609	.020	−.618	−.246	.097	−.611	−.328	.213
Locus of control (6–14)	8.589	8.697	.049	9.749	10.094	.140	9.532	8.816	.057	10.230	8.637	.174	9.742	9.070	.245
Delinquency (0/1)	.671	.824	.118	.842	.798	.157	.848	.880	.011	.803	.854	.090	.846	1.000	.407
Maternal covariates, child age 0:															
Age	22.100	22.230	.183	20.697	23.104	.026	20.058	20.759	.335	20.353	22.448	.091	19.611	19.708	.185
Number of children	1.699	1.430	.001	1.873	1.867	.018	1.954	1.682	.058	1.573	2.362	.262	1.410	1.463	.064
Marital or cohabiting partner (0/1)	.909	.772	.024	.586	.681	.005	.161	.227	.175	.227	.000	.038	.023	.000	.084
High school completion (0/1)	.780	.778	.192	.654	.792	.049	.655	.594	.073	.727	.824	.007	.888	.837	.213

TABLE B1 (*Continued*)

	Stratum 1 $p = [0–.05)$			Stratum 2 $p = [.05–.10)$			Stratum 3 $p = [.10–.15)$			Stratum 4 $p = [.15–.20)$			Stratum 5 $p = [.20–1)$		
	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B	$E(X) \mid d = 0$	$E(X) \mid d = 1$	B
College completion (0/1)	.053	.013	.390	.000	.000	.000	.000	.000	.000	.000	.000	.000	.000	.000	.000
Region (0/1)	.490	.445	.054	.352	.339	.069	.526	.274	.045	.209	.215	.033	.252	.283	.239
Employer tenure (weeks)	15.113	11.668	.101	9.934	9.276	.085	6.470	6.904	.089	4.011	13.331	.240	8.514	9.539	.001
Full-time employment (0/1)	.995	.963	.031	1.000	1.000	.010	1.000	1.000	.000	1.000	1.000	.000	1.000	1.000	.000
Manufacturing worker (0/1)	.067	.019	.048	.185	.064	.020	.122	.169	.034	.076	.312	.282	.424	.360	.155
Trade worker (0/1)	.186	.186	.026	.143	.164	.007	.114	.166	.103	.059	.174	.152	.075	.068	.188
Child outcomes in young adulthood:															
High school completion (0/1)	.802	.835		.664	.594		.733	.666		.837	.672		.845	.857	
Some college attendance (0/1)	.516	.434		.313	.168		.260	.480		.346	.305		.478	.313	
College completion (0/1)	.237	.198		.084	.123		.098	.237		.095	.233		.125	.059	
CESD ages 20–24 (0–1)	.216	.313		.238	.325		.283	.328		.207	.234		.289	.288	
CESD ages 25–29 (0–1)	.198	.206		.198	.348		.225	.233		.222	.286		.319	.281	
n	2,369	172		1,555	223		518	79		294	85		309	93	

NOTE.—Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and who were at least 19 years old in 2010 ($n = 5{,}697$). $E(X) \mid d = 0$ indicates the mean of X for children whose mothers were not displaced, and $E(X) \mid d = 1$ indicates the mean of X for children whose mothers were displaced. B is the standardized difference metric between the treated and the control groups for X.

APPENDIX C

TABLE C1
COVARIATE MEANS FOR CHILDREN OF DISPLACED MOTHERS BY MULTISTATE DISPLACEMENT INDICATORS

	Mother Displaced			
	Recessionary Period	Nonrecessionary Period	LMA with High Unemployment	LMA without High Unemployment
Maternal time-invariant covariates:				
Black (0/1)	.360	.394	.411	.382
Hispanic (0/1)	.111	.056	.172	.035
(Grand)mother's years of education	10.277	10.702	9.975	10.763
(Grand)parents' intact family age 14 (0/1)	.793	.613	.530	.684
(Grand)parents' southern residence (0/1)	.392	.391	.292	.434
Mental ability (ASVAB; −3 to 3)	−.181	−.430	−.331	−.367
Locus of control (6–14)	8.811	9.265	8.871	9.271
Delinquency (0/1)	.731	.888	.896	.827
Maternal covariates, child age 0:				
Age	22.923	21.869	20.860	22.522
Number of children	1.681	1.734	1.505	1.768
Marital or cohabiting partner (0/1)	.549	.510	.499	.525
High school completion (0/1)	.859	.749	.880	.744

TABLE C1 (*Continued*)

	Mother Displaced			
	Recessionary Period	Nonrecessionary Period	LMA with High Unemployment	LMA without High Unemployment
College completion (0/1)	.000	.006	.000	.006
Region (0/1)	.349	.357	.402	.335
Employer tenure (weeks)	8.622	10.647	18.874	7.835
Full-time employment (0/1)	1.000	.982	1.000	.982
Manufacturing worker (0/1)	.143	.103	.205	.091
Trade worker (0/1)	.181	.153	.120	.177
Child outcomes in young adulthood:				
High school completion (0/1)	.818	.723	.727	.748
Some college attendance (0/1)	.306	.352	.393	.336
College completion (0/1)	.072	.168	.235	.122
CESD ages 20–24 (0–1)	.244	.328	.225	.322
CESD ages 25–29 (0–1)	.197	.282	.200	.270
n	203	507	161	491

NOTE.—Sample restricted to children whose mothers were not continuously married and living with a spouse and for whom we have data on job separation, and who were at least 19 years old in 2010 ($n = 5{,}697$). Recessionary periods include 1990–92, 2001–2, and 2008–10; nonrecessionary periods include the remaining periods. High-unemployment areas include any displacement that occurred when the respondent lived in a labor market area (LMA) with at least 9% unemployment; without-high-unemployment areas include those with less than 9% unemployment.

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