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NEIGHBORHOOD POVERTY AND NONMARITAL FERTILITY: SPATIAL AND TEMPORAL DIMENSIONS

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Abstract

Data from 4,855 respondents to the Panel Study of Income Dynamics were used to examine spatial and temporal dimensions of the effect of neighborhood poverty on teenage premarital childbearing. Although high poverty in the immediate neighborhood increased the risk of becoming an unmarried parent, high poverty in surrounding neighborhoods reduced this risk. The effect of local neighborhood poverty was especially pronounced when surrounding neighborhoods were economically advantaged. Measuring exposure to neighborhood poverty over the childhood life course yielded stronger effects than measuring exposure at a single age. Neither racial differences in the level of poverty in proximate neighborhoods nor racial differences in neighborhood poverty over the childhood life course explained the racial difference in nonmarital fertility.

Keywords

Geographic Proximity; Neighborhood; Nonmarital Parenting; Poverty; Race

High rates of nonmarital fertility in the United States (Wu, 2008), coupled with its significant individual and societal costs (U.S. Department of Health and Human Services, 1995), continue to direct scholarly attention to the causes of childbearing outside of marriage (Wu & Wolfe, 2001). Nonmarital childbearing during the teenage years is a particular concern. Although personal, family, and policy determinants are the focus of much of this research, a growing literature has examined the impact of areal—particularly neighborhood—characteristics on teenage pregnancy and out-of-wedlock childbearing (e.g., Brooks-Gunn, Duncan, Klebanov, & Sealand, 1993;Crane, 1991;Harding, 2003). These studies have generated inconsistent results. Some studies found that neighborhood poverty increases the risk of nonmarital childbearing (e.g., Billy and Moore 1992;South & Crowder, 1999;Sucoff & Upchurch, 1998), but other studies failed to find such an effect (e.g., Galster, Marcotte, Mandell, Wolman, & Augustine, 2007;Ginther, Haveman, & Wolfe, 2000;Thornberry, Smith, & Howard 1997).

Two aspects of the typical research design used to identify the influence of neighborhood poverty on nonmarital childbearing potentially limit the utility of most studies on the topic and may help account for these inconsistent findings. First, virtually all studies in this genre considered only the influence of socioeconomic conditions in the immediate neighborhood of residence. But neighborhoods are embedded in a larger system of communities, and it is possible that characteristics of nearby areas may influence the behavior of individuals in a

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given neighborhood (Dietz, 2002; Sampson, 2008). Second, most studies of neighborhood effects on fertility behavior measured neighborhood conditions at a single point during childhood or early adolescence (e.g., Brooks-Gunn et al., 1993; Sucoff & Upchurch, 1998). But measuring neighborhood characteristics at a single point in time ignores the fact that families frequently move into and out of poor neighborhoods (South & Crowder, 1997), and thus focusing exclusively on neighborhood conditions experienced at a single age, rather than throughout the entire childhood life course, may mischaracterize how neighborhood poverty affects fertility behavior in late adolescence and young adulthood (Clampet-Lundquist & Massey, 2008).

This analysis addresses these possible limitations by exploring the spatial and temporal dimensions of the effect of neighborhood poverty on the risk of becoming an unmarried teenage parent. We used spatially-referenced, longitudinal data from the Panel Study of Income Dynamics (PSID) for a cohort of individuals born between 1968 and 1985. We explored the *spatial* dimension of neighborhood effects by using techniques of spatial data analysis to examine how the poverty rate in “extralocal” neighborhoods–-those neighborhoods that surround an individual’s neighborhood of residence—affects the likelihood that PSID sample members bear or father a child outside of marriage. We explored the *temporal* dimension of neighborhood effects by considering how exposure to poor neighborhoods throughout the entire childhood life course affects the likelihood of becoming an unmarried parent.

THEORETICAL BACKGROUND AND HYPOTHESES

Theories of neighborhood effects suggest several reasons why growing up in a poor neighborhood might influence adolescent and young adult behavior, including premarital childbearing (Jencks & Mayer, 1990). Some of these models imply that exposure to poor neighbors impairs children’s social development, but others suggest that exposure to socioeconomically disadvantaged neighbors generates positive outcomes for children. Epidemic or contagion models emphasize the role of peers (Crane, 1991). According to this model, growing up in a poor neighborhood increases the risk of bearing children outside of marriage because in poor communities friends and peers hold pro-childbearing norms, encourage risky sexual activity, and inhibit the use of effective contraception (South & Baumer, 2000). Collective socialization models of neighborhood effects emphasize the influence of nonparental adults on youth fertility behavior. Poor adults fail to provide successful role models for neighborhood children that would discourage early and/or unmarried childbearing (Wilson, 1987). Poor neighborhoods may also lack sufficient levels of monitoring and supervision to deter premarital sexual activity. Institutional models emphasize the importance of schools, churches, and other institutions in providing opportunities for conventional behavior. In poor neighborhoods the relative absence of quality educational and employment opportunities for youth may lead to comparatively fewer social and economic costs to unmarried childbearing (Jencks & Mayer, 1990). The absence of marriageable partners may raise the risk of premarital childbearing by reducing marriage rates and thus increasing the duration of exposure to the risk of unmarried parenthood.

Epidemic, collective socialization, and institutional models of neighborhood effects focus on the socioeconomic characteristics of neighbors and institutions in the immediate neighborhood of residence. But these models may also have implications for the possible impact of extralocal “neighbors”—peers, adults, and institutions outside of the focal neighborhood but nonetheless in close spatial proximity. One possibility is that the level of poverty in surrounding neighborhoods replicates the detrimental effect of poverty in the immediate neighborhood on adolescents’ risk of becoming an unmarried parent. Poverty and

associated disadvantaged conditions in extralocal neighborhoods may spill over into the focal neighborhood. Further, to the extent that social networks extend beyond the boundaries of the immediate neighborhood of residence, peers and adults in surrounding areas are likely to provide important sources of socialization for adolescents. Because contiguous neighborhoods generally exhibit similar poverty rates (i.e., they exhibit positive spatial correlation), it is possible that some of the observed influence of poverty in the immediate neighborhood on nonmarital fertility is instead attributable to the effect of poverty in adjacent or otherwise proximate neighborhoods (Morenoff, 2003; Sampson, Morenoff, & Earls, 1999).

Other theoretical models of neighborhood effects on children's behavior suggest that exposure to poor neighbors may have beneficial outcomes for children or, equivalently, that there are costs to being exposed to nonpoor or otherwise advantaged neighbors. Relative deprivation models maintain that individuals, including children, assess their own social and economic standing in comparison to those around them (Turley, 2002). Exposure to economically advantaged neighbors heightens perceptions of deprivation and diminished social and economic opportunities, rendering youth less likely to postpone childbearing until after marriage. Cultural conflict theories also suggest that disadvantaged neighbors may be a benefit, and advantaged neighbors a liability, for adolescents' social development. Jencks and Mayer (1990) suggested that deviant subcultures, including those that value early and risky sexual activity and childbearing outside of marriage, may be more likely to develop in settings where failure and success are distributed unevenly. According to this argument, greater exposure to successful advantaged neighbors creates a subculture that devalues normatively prescribed behaviors such as childbearing only within marital unions. Resource competition theories suggest that, when surrounded by more advantaged peers, children—perhaps especially disadvantaged children—will lose out in the contest for neighborhood-based resources, such as better schools, better teachers, and employment opportunities, that might deter the transition to unmarried parenthood.

Relative deprivation, cultural conflict, and resource competition processes may be particularly important when considering the influence of extralocal neighbors. Institutions and peers in these proximate neighborhoods may interact frequently enough with youth in a focal neighborhood to serve as comparison groups for perceptions of relative deprivation, to foment deviant, pro-childbearing subcultures, and to create competition for resources, but at the same time be distant enough to avoid transmitting attitudes and norms conducive to nonmarital fertility. In particular, when poor neighborhoods are surrounded by comparatively better-off neighborhoods, these more affluent communities may siphon off resources, such as higher-quality teachers and more generous school funding, that would otherwise serve to reduce nonmarital fertility in the focal community. In contrast to the neighborhood spillover hypothesis advanced above, this reasoning suggests that the poverty rate in extralocal neighborhoods will be *inversely* related to nonmarital childbearing in the focal neighborhood. Moreover, if this hypothesis holds, then it is likely that the level of poverty in extralocal neighborhoods may suppress an effect of local neighborhood poverty on nonmarital childbearing, given the positive spatial correlation of neighborhood poverty rates.

Socioeconomic conditions in extralocal neighborhoods might also moderate the effect of poverty in the immediate neighborhood. Two contrasting scenarios are possible. On one hand, high levels of poverty in extralocal neighborhoods could reinforce the detrimental impact of poverty in the local neighborhood on the risk of becoming an unmarried parent. When poor neighborhoods are surrounded by other poor neighborhoods, youth may be especially isolated from individuals and institutions that might provide employment and educational opportunities, and mainstream cultural values and role models, adverse to

nonmarital childbearing. On the other hand, relative deprivation theory suggests that local neighborhood poverty and extralocal neighborhood poverty may interact negatively in affecting the risk of nonmarital childbearing. Youth residing in poor neighborhoods may perceive their social and economic opportunities as especially restricted when the socioeconomic conditions in their own neighborhood compare unfavorably to those in surrounding areas. Under this scenario, the influence of poverty in the immediate neighborhood on nonmarital childbearing will be weaker, rather than stronger, when the focal neighborhood is surrounded by equally poor neighborhoods.

Addressing the possible effects of poverty in extralocal neighborhoods may have implications for explaining the pronounced racial difference in teenage nonmarital childbearing. Prior studies have generated somewhat equivocal conclusions regarding the degree to which racial differences in the socioeconomic status of residential neighborhoods can explain racial differences in fertility and sexual behavior (Brewster, 1994; Browning, Leventhal, & Brooks-Gunn, 2004; South & Baumer 2000). But these studies rarely incorporated measures of poverty in surrounding neighborhoods. Because the geographic concentration of poverty is more pronounced among Blacks than among Whites (Massey & Denton, 1993), Blacks are not only more likely than Whites to encounter peers who support nonmarital childbearing, economically-disadvantaged adults, and weak institutions within their immediate neighborhood, they are also more likely than Whites to experience similarly unfavorable conditions in adjacent communities. If high poverty in extralocal neighborhood exacerbates the risk of nonmarital childbearing in focal neighborhoods, Blacks' greater exposure to poverty in surrounding communities may help to explain the racial difference in nonmarital fertility even after controlling for the level of poverty in the immediate neighborhood.

Exposure to Neighborhood Poverty over the Childhood Life Course

Another possible limitation to the research designs most often used to examine the effects of neighborhood poverty on nonmarital fertility is the almost exclusive reliance on point-in-time measures of neighborhood poverty. Given considerable migration between poor and nonpoor neighborhoods, studies that measured neighborhood poverty at a single point in time (Sucoff & Upchurch, 1998), or even averaged over a few childhood years (Harding, 2003), may have failed to capture meaningful levels of exposure to neighborhood poverty. Neighborhood poverty may only matter for children who have lived in their neighborhood for a substantial period of time (Turley, 2003).

Some studies suggested that single-year measures of neighborhood characteristics are reasonable proxies for measures of children's long-run neighborhood environments (Jackson & Mare, 2007; Kunz, Page, & Solon, 2003), but these studies relied on generally short time frames for these comparisons. Over the entire childhood life course, there is likely to be greater intraindividual variation in neighborhood poverty both because children's families will have had more time to move to a different neighborhood with a different poverty rate and because neighborhoods themselves will have had a longer time to experience a change in their level of poverty. We explored this possible limitation of studies of neighborhood effects on nonmarital childbearing by computing a cumulative level of neighborhood poverty experienced by children throughout their life course. We then compared the effect on nonmarital childbearing of neighborhood poverty measured over the childhood life course with the more conventional strategy of measuring neighborhood poverty at a single age.

Explicitly modeling the impact of the exposure to neighborhood poverty over the childhood life course may also have implications for explaining the pronounced racial difference in the risk of becoming an unwed parent. Black children spend a significantly greater portion of

their childhood in poor neighborhoods compared to their White counterparts, and racial differences in the cumulative exposure to poor neighborhoods over the childhood life course are greater than racial differences at any single age (Timberlake, 2007). Consequently, incorporating racial differences in exposure to local neighborhood poverty throughout the childhood life course may help to explain that part of the Black-White difference in premarital childbearing that cannot be explained by established individual-level and family-level predictors or by the level of neighborhood poverty experienced at any single age.

Other Predictors of Nonmarital Childbearing

Attempting to infer a causal effect of local and extralocal neighborhood poverty on the risk of nonmarital fertility necessitates consideration of other predictors of out-of-wedlock childbearing, especially factors that might be correlated with neighborhood socioeconomic status. Chief among these are the economic resources and composition of youths' family of origin. Greater parental economic resources are likely to reduce the risk of premarital fertility by encouraging youth to pursue educational and employment opportunities that conflict with early childbearing (Duncan, Yeung, Brooks-Gunn, & Smith, 1998). Growing up in a single-parent family encourages premarital childbearing partly because of the low and unstable incomes of single-parent families (McLanahan & Sandefur, 1994). Residential mobility has also been shown to increase the risk of premarital childbearing, perhaps because geographic relocation increases the risk of early sexual activity (South, Haynie, & Bose 2005). Residents of metropolitan areas may face higher risks of premarital childbearing than residents of nonmetropolitan communities. The risk and timing of becoming an unmarried parent are also likely to vary by race, age, and birth cohort.

METHOD

Our main source of data for this analysis is the Panel Study of Income Dynamics (PSID), a longitudinal survey of U.S. residents and their families (Hill, 1992). Members of the initial panel of approximately 5,000 families were interviewed annually until 1997 and biennially thereafter. New families have been added to the panel as children and other members of original panel families have formed their own households. The PSID has been used widely in studies of neighborhood effects on adolescent and young adult behavior, including teenage pregnancy and premarital childbearing (e.g., Brooks-Gunn et al., 1993; Harding, 2003; South & Crowder, 1999). We appended tract-level data from four decennial U.S. censuses (1970 to 2000) to the PSID sample members' individual records to capture the level of poverty in the immediate and geographically-proximate neighborhoods that they experienced throughout their childhood life course.

Sample selection

We initially selected all black and white PSID participants who were born into a PSID panel family between 1968 and 1985 (n = 6,069). (Members of other racial and ethnic groups were either represented in too few numbers or were not followed long enough by the PSID to warrant inclusion in this analysis.) We omitted 18% of respondents who lacked required data—collected many years later—on the timing of marriage and childbearing (n = 815), and following Wagmiller, Lennon, Kuang, Alberti, and Aber (2006) we further omitted 5% of respondents for whom valid information was not collected for at least seven years when the respondents were between the ages of 0 and 13 (n = 301). Finally, we omitted 2% of respondents (n = 98) who were not in the sample at age 14 when respondents entered the risk set for becoming an unmarried parent. These selections resulted in a sample of 4,855 individuals, of whom 12.2% reported becoming an unmarried parent between the ages of 14 and 20. Compared to the PSID sample members who are excluded from the analysis, the PSID sample members who are retained in the analysis were less likely to be black (47.1%

vs. 53.10%), less likely to have been born into a female-headed family (17.8% vs. 20.9%), and less likely to have been born into a poor family (25.9% vs. 33.9%).

Measuring Local and Extralocal Neighborhood Poverty

Following most prior research in this area, we used census tracts as our approximation of neighborhoods. Tract-level census data were drawn from the Neighborhood Change Data Base (NCDB), in which data from earlier censuses (1970, 1980, and 1990) have been normalized to 2000 tract boundaries, allowing us to produce consistent measures of neighborhood poverty over time (GeoLytics, 2008). Although neighborhood socioeconomic status can be measured by a variety of indicators (Duncan, Connell, & Klebanov, 1997), we used the official poverty rate which, unlike multi-item scales, has a straightforward interpretation (Galster et al., 2007). We took the census tract poverty rate at each census year and used linear interpolation to estimate values for noncensal years. We then attached the poverty rate to the records of the PSID respondents according to their tract of residence at each year they were in the sample.

To measure the level of poverty in extralocal neighborhoods, we computed a spatially-weighted average level of poverty in census tracts surrounding each PSID respondent's tract of residence. This variable is based on the application of spatial weights to the poverty rate in these extralocal areas, with the magnitude of the spatial weights corresponding to the presumed influence of conditions in each surrounding tract on behaviors of individuals in the tract of residence. Following Downey's (2006) argument that spatial dependence tends to decline with distance, we employed a spatial-weighting strategy in which the influence of poverty in an extralocal tract on nonmarital childbearing is assumed to be inversely related to the distance of the extralocal tract from the individual's tract of residence. Specifically, under this distance-decay strategy we developed a row-standardized matrix of spatial weights defined as $w_{ij} = 1/d_{ij}$ where d_{ij} is the geographic distance between the centroid of the tract of residence (i) and the centroid of the extralocal tract, j. Given the implausibility that the poverty rate of every tract in the nation directly affects the decisions of residents in all other tracts, we constrained to zero the influence of tracts that are more than 100 miles away from the focal tract, but even without this constraint, spatial weights based on inverse distance approach zero – representing an absence of influence – beyond distances of about 10 miles. As with the poverty rate in the neighborhood of immediate residence, we estimated values for this average "extralocal poverty rate" from the NCDB from the decennial censuses and attached them to the records of each PSID respondent at each year they are in the sample.

Measuring Exposure to Neighborhood Poverty over the Childhood Life Course

We computed three measures of the PSID respondents' exposure to poverty in their immediate neighborhood and in their extralocal neighborhoods. First, we employed the common strategy of measuring the local poverty rate experienced at age 14 (e.g., Brooks-Gunn et al., 1993). Second, we computed the average tract poverty rate that the respondents experienced between birth and the year preceding each age that they were at risk of becoming an unmarried parent. We focused on the level of tract poverty in the year preceding each age because these are the conditions relevant to the potential conception of a birth occurring in the subsequent year. Third, we computed the average level of the extralocal poverty rate experienced between birth and the year preceding each age respondents were at risk of becoming an unmarried parent, using the spatial weighting strategy described above. Not surprisingly, the correlation between the tract poverty rate of the local neighborhood (averaged over the childhood years) and the tract poverty rate of extralocal neighborhoods (also averaged over the childhood years) was moderately high (r = .697), but diagnostic checks did not reveal evidence of multicollinearity problems.

Control variables

We controlled for several individual and family characteristics that might be related to both neighborhood poverty and the likelihood of having a child outside of marriage. The duration dependence of the risk of nonmarital childbearing was captured by dummy variables for respondents' age. Respondent's race is a dummy variable scored 0 for Whites and 1 for Blacks. Because men are likely to underreport fathering children outside of marriage (Rendall, Clarke, Peters, Ranjit, & Verropoulou, 1999), we included a dummy variable for respondent's gender, scored 0 for men and 1 for women. Secular trends in nonmarital childbearing were captured by a continuous variable for year of birth. Younger individuals thus received higher scores than older individuals on this measure of birth cohort. Family economic status was measured by the cumulative proportion of years between birth and the observation year that the respondent's family had an income below the poverty level. Parental educational attainment was measured by a dummy variable scored 1 for sample members whose family head (most often the father) had completed college by the time the sample member was age 14. Childhood family structure was measured by the cumulative proportion of childhood years in which the individual's family was headed by a female (most often the mother). Residential mobility was measured by the proportion of childhood years in which the respondent moved between census tracts from one year to the next. The models included a dummy variable indicating whether, at the beginning of each annual risk period, the respondent was residing in a metropolitan area.

Analytical Strategy

To examine the impact of local and extralocal neighborhood poverty on the timing of the transition to unmarried parenthood, we estimated a series of discrete-time event-history regression models (Allison, 1984). In these models each respondent's childbearing history (beginning at age 14) was segmented into a series of person-year observations. The 4,855 sample members contributed 24,462 person-years to the analysis. The dependent variable was a binary variable indicating whether a (first) nonmarital birth occurred during the interval, and the time-varying independent variables, including the cumulative measures of local and extralocal neighborhood poverty, captured characteristics of the respondents in the year preceding the interval. Respondents who "survived" until age 20 without having become an unmarried parent, as well as those who married prior to becoming a parent or who left the panel without experiencing unmarried parenthood, were treated as censored observations at the time of their exit from the risk set.

We estimated models with both point-in-time and childhood-averaged (i.e., cumulative) measures of exposure to neighborhood poverty to determine the sensitivity of findings to the life course time frame captured by these measures. We included in the models spatially-lagged measures of neighborhood poverty to ascertain whether the likelihood of having a child outside of marriage responds to the poverty rate in extralocal neighborhoods, net of any response to the level of poverty in the immediate neighborhood. We estimated models with and without the spatially-lagged and childhood exposure measures of neighborhood poverty to determine their ability to explain the pronounced racial difference in nonmarital childbearing. And, we incorporated the relevant interaction term in these models to determine whether the impact of poverty in the immediate neighborhood is moderated by the level of poverty in extralocal neighborhoods. Tests of statistical significance were based on robust standard errors that account for the clustering of observations within census tracts.

RESULTS

Table 1 presents descriptive statistics from the event-history person-year file for all variables used in the analysis. In the typical observation year, about 2.4% of the at-risk respondents

became an unmarried parent. The average age of the sample members at the beginning of each person-year observation was 16.2 years. The sample was 45% Black (a consequence of the PSID's oversampling of poor families) and 48% female. The "average" sample member was born in 1977. On average, respondents spent about 24% of their childhood years leading up to the observation period in a poor family. Eighteen percent of the sample members came from a family in which the head completed college. The sample members spent about 24% of their childhood years in a female-headed family and moved annually from one tract to another 16% of their childhood years. Almost 80% of the person-years originated in a metropolitan area.

The mean neighborhood poverty rate for the sample members at age 14 was 16.7%, very similar to the mean of the average neighborhood poverty rate experienced during the childhood years (17.2%). The mean spatially-lagged measure of the poverty rate, a weighted average of poverty in surrounding neighborhoods, was 14.8%.

Table 2 presents the results of a series of discrete-time logistic regression event-history models examining how exposure to poor neighbors in both the immediate neighborhood and extralocal neighborhoods influences the likelihood of experiencing nonmarital parenthood. Model 1 is a baseline model that includes as predictors only the individual demographic and family background variables. For the most part these potential determinants of nonmarital childbearing operated as expected.

The annual risk of becoming an unmarried parent over the teenage years increased with age. Blacks were significantly more likely than Whites, and women were significantly more likely than men, to report having a nonmarital birth. Net of the other predictors, there was a significant downward secular trend in teenage nonmarital fertility over the study period. The proportion of childhood years spent in a poor family and the proportion of years spent in a female-headed family were positively and significantly related to the odds of having a nonmarital birth. Childhood residential mobility was also positively and significantly related to the risk of becoming an unmarried parent. Respondents whose family head completed college were significantly less likely than children of less educated family heads to have had an out-of-wedlock birth. Net of the effects of the other predictors, however, the difference between residents of metropolitan and nonmetropolitan areas in the risk of becoming an unmarried parent was not statistically significant.

Model 2 of Table 2 adds the local neighborhood poverty rate experienced at age 14 as an explanatory variable. This model is thus typical of the most common strategy for detecting the influence of neighborhood poverty on the risk of teenage nonmarital childbearing (e.g., Brooks-Gunn et al., 1993). The coefficient for the local neighborhood poverty rate at age 14 was not statistically significant ($b = .000$, $p > .05$).

Model 3 substitutes for the local neighborhood poverty rate at age 14 the average local neighborhood poverty rate experienced during childhood. The coefficient for this cumulative measure of poverty in the local neighborhood was also statistically nonsignificant ($b = .005$, $p > .05$). Moreover, controlling for the level of neighborhood poverty experienced throughout childhood did little to explain the pronounced racial difference in the risk of becoming an unmarried parent. In Model 3, the coefficient for Black race (b = 1.004, OR = 2.730) differed negligibly from the corresponding coefficient in Model 1 (b = 1.059, OR = 2.884).

Model 4 of Table 2 adds to Model 3 the measure of extralocal neighborhood poverty averaged over the childhood years. This explanatory variable taps the level of poverty in neighborhoods surrounding the respondents' tract of residence at each age, with the poverty rate of geographically closer tracts given greater weight than the poverty rate of more distant

tracts. The coefficient for this average extralocal neighborhood poverty rate was negative and statistically significant (b = −.034, $p < .05$). A one standard deviation difference in the average extralocal neighborhood poverty rate translates into a 14% reduction in the risk of experiencing nonmarital teen parenthood [$(.14 = 1 - e^{(-.034)(4.575)}$], and a two standard deviation difference in the average extralocal neighborhood poverty rate translates into a 27% reduction in the risk of becoming an unmarried teen parent [$(.27 = 1 - e^{(-.034)(4.575)(2)}$]. Net of the effects of the other predictors (including the average poverty rate over the childhood years in the local neighborhood), higher levels of poverty in surrounding neighborhoods were associated with lower risks of nonmarital parenthood. Put another way, youth were more likely to have a child outside of marriage when (all else equal) their neighborhoods are enveloped not by poor residents, but rather by nonpoor people. This inverse (net) effect on nonmarital parenthood of the average extralocal neighborhood poverty rate is consistent with theories of neighborhood effects that emphasize relative deprivation, cultural conflict, and resource competition. Yet, racial differences in exposure to extralocal poverty did not explain the pronounced racial difference in the risk of becoming an unmarried parent. The coefficient for Black race actually increased slightly from Model 3 (b = 1.004, OR = 2.730) and Model 4 (b = 1.074, OR = 2.926).

Importantly, controlling for the effect of extralocal neighborhood poverty (averaged over the childhood years) caused the coefficient for local neighborhood poverty (also averaged over the childhood years) to become statistically significant ($b = .012$, $p < .05$). Given the parameter estimates in Model 4, a one standard deviation difference in the neighborhood poverty rate averaged across the childhood years translates into a 16% increase in the odds of becoming an unmarried teen parent [$(1.163 = e^{(.012)(12.610)}$], and a two standard deviation difference in the neighborhood poverty rate averaged across the childhood years translates into a 35% increase in the odds of experiencing teenage nonmarital parenthood [$(1.353 = e^{(.012)(12.610)(2)}$]. Thus, the inverse effect of extralocal neighborhood poverty on the risk of becoming an unmarried teenage parent suppressed the facilitative effect of local neighborhood poverty. Statistically, this suppression of the effect of local neighborhood poverty on the risk of nonmarital parenthood is generated by positive spatial correlation of the neighborhood poverty rate combined with the net inverse effect of the average extralocal neighborhood poverty rate. In substantive terms, this result suggests that studies that have ignored the influence of poverty in surrounding neighborhoods may have nontrivially underestimated the association between local neighborhood poverty and teenage premarital childbearing.

In Model 5 of Table 2 we re-evaluated the effect of the neighborhood poverty rate experienced at age 14 relative to the corresponding effect of neighborhood poverty averaged over the childhood years. Comparing Model 5 with Model 4, it is apparent that, when the influence of extralocal neighborhood poverty (averaged over the childhood years) is controlled, it makes a difference when during the childhood life course exposure to local neighborhood poverty is measured. When measured at age 14 (Model 5), the effect of local neighborhood poverty on the risk of becoming an unmarried parent was weak and statistically nonsignificant ($b = .003$, $p > .05$). But when local neighborhood poverty is measured by averaging its values over the childhood years (Model 4), the coefficient was much larger and statistically significant ($b = .012$, $p < .05$).

Model 6 of Table 2 adds the product term representing the interaction between the poverty rate in the local neighborhood (averaged over the childhood years) and the spatially-lagged poverty rate in extralocal neighborhoods (also averaged over the childhood years). The coefficient for this product term was negative and statistically significant ($b = -.002$, $p < .05$). The facilitating impact of high poverty in the immediate neighborhood of residence on

the risk of nonmarital parenthood weakens as the poverty rate in surrounding neighborhoods increases. Put another way, the detrimental impact of poverty in the immediate neighborhood of residence on the risk of nonmarital parenthood is stronger when these neighborhoods are surrounded by many nonpoor (rather than poor) residents. This finding is consistent with the hypothesis, drawn from relative deprivation theory, that youth residing in poor neighborhoods will perceive their opportunities as particularly restricted when the socioeconomic conditions in their own neighborhood compare unfavorably to those in surrounding communities.

Figure 1 illustrates how the effect of the local neighborhood poverty rate (averaged over the childhood years) on the annual risk of becoming an unmarried parent varied by the poverty rate in extralocal neighborhoods (also averaged over the childhood years). This figure shows the predicted annual probabilities of having a child outside of marriage at varying levels of the local neighborhood poverty rate under three different conditions: when the average poverty rate in extralocal areas was at the 10^{th}, 50^{th}, and 90^{th} percentiles. These probabilities were derived from Model 6 of Table 2, holding all other variables constant at their sample means. As shown in the dark solid line of this figure, when the poverty rate in extralocal neighborhood was very high (at the 90^{th} percentile), the effect of the local poverty rate on nonmarital parenthood was weak and inverse. The predicted risk of having a child outside of marriage actually decreased slightly from a local poverty rate of 0 (predicted probability = .013) to a local poverty rate of 40%, which is about the 95^{th} percentile (predicted probability = .005).

In contrast, as shown in the dotted line of Figure 1, when the extralocal poverty rate was at the 50^{th} percentile (and hence fewer of the extralocal neighbors were poor and more were nonpoor), the annual risk of experiencing nonmarital parenthood increased with increases in the local poverty rate. Under this scenario, the annual risk of having a child outside of marriage rose from about .008 when the local poverty rate was 0 to about .016 when the local poverty rate was 40%. And, as shown in the light grey line of Figure 1, when the extralocal poverty rate was very low, the risk of nonmarital childbearing increased dramatically with increases in the local poverty rate. When the level of extralocal poverty was at the 10^{th} percentile, with only about 4% of extralocal neighbors poor (and thus over 96% of these neighbors were nonpoor), the annual probability of experiencing nonmarital parenthood increased from about .007 when the local poverty rate was 0 to .031 when the local poverty rate was 40%. In sum, being surrounded by nonpoor residents in extralocal neighborhoods sharply increased the impact of local neighborhood poverty on the risk of becoming an unmarried mother or father during the teenage years.

Supplemental Analyses

We performed several supplemental analyses to assess the robustness of our results. First, because census tracts in metropolitan areas differ in geographic size between tracts in nonmetropolitan areas, we examined whether the associations between local and extralocal neighborhood poverty rates and the risk of becoming an unmarried parent vary between residents of metropolitan and nonmetropolitan areas. Regression models that incorporated product terms representing the interaction between metropolitan (versus nonmetropolitan) location and the measures of neighborhood poverty revealed no evidence that these associations varied between metropolitan and metropolitan areas. Second, using similar procedures, we examined whether the effects of local and extralocal neighborhood poverty varied by respondents' race or age. We found no evidence that this is the case; coefficients for the relevant interaction terms were consistently nonsignificant. Third, we examined whether the effect of extralocal neighborhood poverty on the risk of becoming an unmarried parent might be reflecting differences in racial composition between neighborhoods that are surrounded by poor neighborhoods compared to neighborhoods that are surrounded by

nonpoor neighborhoods. But the effect of extralocal neighborhood poverty remained negative and statistically significant even controlling for the percentage of the local tract population that is Black. Fourth, we compared the inverse-distance weighting strategy for measuring extralocal neighborhood poverty to several alternatives, including an adjacent-tracts approach in which only census tracts that share a border with the focal tract are considered influential. The adjacent-tracts weighting strategy yielded weaker findings than strategies that incorporate distance-weighted information from a larger set of surrounding tracts. These findings suggest that characteristics of tracts beyond adjacent tracts impact premarital childbearing rates in the focal tract but that the influence of these extralocal characteristics declines as a function of the distance from the tract of residence.

DISCUSSION

Given the substantial cost of out-of-wedlock childbearing to children, their parents, and society at large, the causes of nonmarital fertility continue to elicit considerable scholarly attention. Although recent studies have sought to locate the causes of nonmarital fertility in community environments, research into neighborhood effects on out-of-wedlock childbearing has been limited both by the failure to consider the influence of neighborhoods outside of, but proximate to, youth’s neighborhood of residence and by the failure to consider children’s exposure to neighborhood poverty over the entire childhood life course. This analysis used geographically-referenced, longitudinal data from the Panel Study of Income Dynamics to address these issues and to expand on prior studies of neighborhood effects on the likelihood of becoming an unmarried mother or father.

We acknowledge several limitations to our study. Because our analysis followed respondents for a long period of time and required data from multiple time points, our effective sample differed somewhat from the full sample of individuals born into the PSID between 1968 and 1985. This sample attrition might have affected our results in unknown ways. We also acknowledge the difficulties inherent in drawing causal inferences about neighborhood effects from nonexperimental designs such as ours. The selection of families into particular types of neighborhoods based on propensities for their children to become unmarried parents represents a possible non-causal explanation for our findings (Harding, 2003). At the same time, however, it is difficult to envision plausible residential decision-making strategies that, all else equal, would lead youth with a high latent propensity toward nonmarital fertility to move to neighborhoods that are surrounded by relatively *nonpoor* neighborhoods. The typical neighborhood selection process, in which youth (and their families) predisposed toward adverse outcomes move to high-poverty neighborhoods, cannot easily explain the *inverse* association between extralocal neighborhood poverty and the risk of becoming an unmarried parent. In this sense, our results underscore Kling, Liebman, and Katz’s (2007) conclusion that if endogenous neighborhood selection accounts for observed associations between neighborhood characteristics and youth outcomes, these selection processes must be quite complex.

Five main conclusions emerge from our analysis. First, we found significant effects of the level of poverty in neighborhoods that surround youths’ neighborhood of residence on the risk of becoming an unmarried teenage parent. Consistent with theories of relative deprivation, cultural conflict, and resource competition, higher levels of poverty in extralocal neighborhoods were associated with *lower* risks of premarital childbearing in the focal neighborhood, net of other determinants. Several overlapping mechanisms could conceivably account for the inverse net effect of extralocal neighborhood poverty on nonmarital parenthood. As suggested by relative deprivation theory, youth may feel especially deprived of opportunities that protect against nonmarital childbearing when their own neighborhoods are surrounded by relatively advantaged communities. Unfavorably

comparing their own social and economic opportunities to those of better-off extralocal neighbors may generate a sense of fatalism that deters the desire to postpone childbearing until after marriage (Harris, Duncan, & Boisjoly, 2002), an explanation broadly congruent with cultural conflict models. Consistent with the resource competition model of neighborhood effects, relatively advantaged proximate neighborhoods may also siphon off resources conducive to postponing parenthood that would otherwise be targeted to the focal neighborhood. In particular, advantaged proximate neighborhoods may divert educational resources—such as better teachers, instructional technologies, parental and school social capital, and funding—and employment opportunities away from the focal neighborhood. In turn, these diminished educational and employment opportunities may encourage nonmarital childbearing among youth in the focal neighborhood by reducing the opportunity costs to unmarried parenthood. Future research might profit from exploring these and other explanations for the higher rates of nonmarital fertility in neighborhoods that are spatially enveloped by economically advantaged communities.

A second key conclusion is that taking into account this inverse effect of extralocal neighborhood poverty substantially increases the observed effect of local neighborhood poverty on the risk of becoming an unwed mother or father. Indeed, we failed to observe a significant effect of local neighborhood poverty on the risk of becoming an unmarried parent when the impact of extralocal neighborhood poverty was not controlled. By failing to consider the role played by the level of poverty in extralocal neighborhoods, prior studies may have underestimated the effect of poverty in the immediate neighborhood on the risk of nonmarital childbearing.

Third, we found that local and extralocal neighborhood poverty interact in affecting nonmarital childbearing. The facilitative effect of local neighborhood poverty is substantially stronger when the neighborhood is surrounded by mostly nonpoor rather than poor residents. Exposure to local neighborhood poverty appeared to have little effect on the risk of nonmarital parenting when surrounding neighborhoods also had high concentrations of poverty. This finding, too, seems at least broadly consistent with relative deprivation theory. Youth residing in poor neighborhoods are likely to perceive their social and economic alternatives to nonmarital childbearing to be even more restricted when the economic status of their own neighborhood compares unfavorably to that in surrounding areas. The combination of high poverty in the local neighborhood with low poverty in proximate neighborhoods generates the highest risks of becoming an unmarried mother or father.

Our fourth conclusion concerned the temporal dimension of the effect of neighborhood poverty on nonmarital fertility. Most prior studies of the effects of neighborhood poverty on nonmarital fertility measure exposure to neighborhood poverty at a single point in the childhood lifecourse, ignoring the fact that families frequently move between poor and nonpoor neighborhoods. After controlling for extralocal neighborhood poverty, we found substantially stronger effects of local neighborhood poverty when exposure was measured as the average level of neighborhood poverty experienced during the childhood years rather than only at age 14. Future research into the temporal dimension of neighborhood poverty's effect on nonmarital fertility might benefit by attending to possible age or lifecourse variation in this effect. For example, it is likely that exposure to poor neighborhoods during the early childbearing years is more consequential for fertility behavior than similar exposure during early childhood.

The fifth and final conclusion to be drawn from our study involves the limited ability of racial differences in exposure to extralocal neighborhood poverty and in exposure to neighborhood poverty over the childhood life course to explain the pronounced racial

difference in nonmarital childbearing. We found that none of the racial difference in the risk of becoming an unmarried parent can be attributed to racial differences in exposure to poverty in surrounding neighborhoods. To be sure, Black youth are more exposed than White youth to high levels of poverty both in their immediate neighborhood and in surrounding communities, but because the poverty rate in extralocal neighborhoods was inversely related to nonmarital parenthood, the racial difference in exposure to extralocal neighborhood poverty tended to suppress, rather than to explain, the racial difference in nonmarital fertility. And, although Black youth are more exposed than White youth to neighborhood poverty over their childhood life course, this racial difference in poverty exposure explained little of the racial difference in the risk of becoming an unmarried parent. Thus, although this research highlights important temporal and spatial dimensions of neighborhood effects, comprehensive explanations for racial differences in nonmarital fertility will need to look beyond both racial differences in the exposure to local neighborhood poverty over the childhood life course and racial differences in exposure to poverty in nearby neighborhoods.

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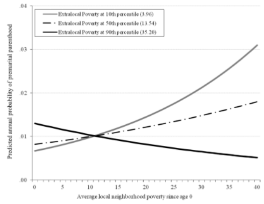

Figure 1.

Predicted Annual Probability of Premarital Parenting by Local and Extralocal Levels of Neighborhood Poverty for Panel Study of Income Dynamics Participants Born between 1968 and 1985

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Table 1

Descriptive Statistics for Variables Used in Event-History Analysis of Premarital Parenting for Panel Study of Income Dynamics Participants Born between 1968 and 1985

Variable	**Definition**	***M***	***SD***
Dependent Variable			
Premarital parenting	Whether R became a parent in year of observation (1=yes)	.024	.154
Independent Variables			
Local Neighborhood Conditions			
Local neighborhood poverty at age 14	Percent of population in poverty in R's tract of residence at age 14	16.673	13.734
Average local neighborhood poverty since age 0	Percent of population in poverty in R's tract of residence averaged from age 0 to year preceding observation	17.197	12.610
Extralocal Neighborhood Conditions			
Average extralocal neighborhood poverty since age 0	Distance-weighted average poverty rate in tracts surrounding R's tract of residence, averaged from age 0 to year preceding observation	14.761	4.575
Micro-level characteristics			
Age	Age of R in year of observation	16.241	1.666
Black	Whether R lived in a household with a Black household head (1=yes)	.452	.498
Female	Whether R is female (1=yes)	.478	.500
Year of birth	Year R was born	1977.083	5.060
Time in poor family	Proportion of years between age 0 and year preceding observation in which R lived in a family with income below poverty threshold	.240	.331
Head completed college	Whether head of R's household at age 14 completed college (1=yes)	.181	.385
Time in female-headed family	Proportion of years between age 0 and year preceding observation in which R lived in a family headed by a single woman	.237	.351
Residential mobility	Proportion of consecutive years between age 0 and year preceding observation in which R's family moved to a different tract	.162	.156
Metropolitan resident	Whether R lived in a census Metropolitan Statistical Area in year of observation	.790	.407

Note: Number of person-year observations = 24,462; Number of individuals = 4,855.

Table 2

Summary of Logistic Regression Analysis of Variables Predicting Premarital Parenting for Panel Study of Income Dynamics Participants Born between 1968 and 1985

	Model 1		Model 2		Model 3		Model 4		Model 5		Model 6	
	B	***SE B***	***B***	***SE B***	***B***	***SE B***	***B***	***SE B***	***B***	***SE B***	***B***	***SE B***
Local Neighborhood Conditions												
Local neighborhood poverty at age 14			.000 (1.000)	.003					.003 (1.011)	.004	.047*** (1.011)	.014
Average local neighborhood poverty since age 0					.005 (1.005)	.004	.012* (1.011)	.005				
Extralocal Neighborhood Conditions												
Average extralocal neighborhood poverty since age 0							−.034* (.966)	.014	−.022 (1.011)	.014	.021 (1.011)	.026
Local-Extralocal Interaction												
Local poverty since age 0 × extralocal poverty since age 0											−.002* (1.011)	.001
Micro-level characteristics												
Age^{a}												
15	.640** (1.897)	.241	.640** (1.897)	.241	.640** (1.897)	.241	.643** (1.903)	.241	.642** (1.900)	.241	.644** (1.905)	.241
16	1.554*** (4.731)	.214	1.554*** (4.731)	.214	1.554*** (4.732)	.214	1.561*** (4.763)	.214	1.557*** (4.746)	.214	1.561*** (4.766)	.214
17	1.811*** (6.118)	.219	1.811*** (6.117)	.219	1.812*** (6.122)	.219	1.821*** (6.176)	.219	1.815*** (6.141)	.219	1.819*** (6.167)	.219
18	2.191*** (8.940)	.211	2.191*** (8.940)	.211	2.193*** (8.958)	.211	2.207*** (9.089)	.211	2.198*** (9.004)	.211	2.205*** (9.075)	.211
19	2.533*** (12.589)	.210	2.533*** (12.589)	.210	2.535*** (12.613)	.210	2.550*** (12.811)	.210	2.541*** (12.686)	.210	2.549*** (12.798)	.210
Black	1.059*** (2.885)	.132	1.059*** (2.883)	.136	1.004*** (2.730)	.141	1.074*** (2.926)	.143	1.125*** (3.079)	.141	.949*** (2.582)	.149
Female	.893*** (2.442)	.088	.893*** (2.442)	.088	.893*** (2.443)	.088	.895*** (2.447)	.088	.891*** (2.437)	.088	.901*** (2.463)	.088
Year of birth	−.021** (.979)	.008	−.021** (.979)	.008	−.021** (.979)	.008	−.017* (.983)	.008	−.018* (.982)	.008	−.017* (.984)	.008
Time in poor family	.749*** (2.115)	.150	.748*** (2.113)	.157	.685*** (1.984)	.157	.729*** (2.073)	.156	.796*** (2.217)	.157	.716*** (2.047)	.155
Head completed college	−1.220*** (.295)	.251	−1.220*** (.295)	.252	−1.209*** (.298)	.252	−1.194*** (.303)	.253	−1.212*** (.297)	.253	−1.151*** (.316)	.253
Time in female-headed family	.277* (1.319)	.134	.277* (1.319)	.136	.259 (1.296)	.136	.214 (1.239)	.138	.253 (1.288)	.138	.253 (1.225)	.138

	Model 1		Model 2		Model 3		Model 4		Model 5		Model 6	
	B	***SE B***	***B***	***SE B***	***B***	***SE B***	***B***	***SE B***	***B***	***SE B***	***B***	***SE B***
Residential mobility	1.139*** (3.124)	.277	1.140*** (3.126)	.276	1.181*** (3.259)	.278	1.152*** (3.167)	.277	1.100*** (3.004)	.279	1.107*** (3.025)	.280
Metropolitan resident	.029 (1.029)	.126	.029 (1.029)	.126	.028 (1.028)	.126	−.034 (.970)	.130	−.009 (.991)	.130	.008 (1.008)	.132
Constant	35.15		35.12		34.70		26.29		29.19		25.16	
χ^2	806.00		809.11		807.13		820.17		821.19		842.87	
Df	13		14		14		15		15		16	

Note: Odds ratios in parentheses. n of person-year observations = 24,462; n of individuals = 4,855; n of premarital births=593.

[a]Reference category for age is age 14.

*p < .05.

**p < .01.

***p < .001.