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The Neighborhood Contribution to Black-White Perinatal Disparities: An Example From Two North Carolina Counties, 1999–2001

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Previous studies of black-white disparities in perinatal outcomes have generally not controlled for both observed and unobserved neighborhood inequalities with models that compare only black and white women living in the same neighborhoods. Using 1999–2001 birth certificate data from 2 counties in North Carolina, the authors employed a hybrid fixed-effects approach to assess the total contribution of neighborhood factors to both absolute and relative racial disparities in low birth weight, preterm birth (PTB), and smallness for gestational age at term. Neighborhood factors made a notable contribution to racial disparities for PTB only, accounting for an additional 15% reduction in crude disparities beyond individual sociodemographic characteristics, which accounted for approximately 40% of racial disparities. The neighborhood contribution was greater for moderate PTB (32–36 weeks' gestation) than for very PTB (<32 weeks' gestation). A neighborhood deprivation index accounted for a smaller percentage of PTB disparities than the hybrid fixed-effects estimates, which suggests that measured socioeconomic deprivation does not account for all health-relevant neighborhood inequalities. Contemporaneous individual-level sociodemographic and neighborhood factors together explained one- to two-thirds of perinatal disparities. To fully explain racial disparities in perinatal outcomes, evaluation of other differential exposures (e.g., racism or wealth) and neighborhood factors across the life course may be necessary.

health status disparities; infant, low birth weight; infant, small for gestational age; minority health; multilevel analysis; premature birth; residence characteristics

Abbreviations: LBW, low birth weight; MPTB, moderately preterm birth; PTB, preterm birth; RR, relative risk; SES, socioeconomic status; SGA, small for gestational age; VPTB, very preterm birth.

Black infants are approximately twice as likely as white infants to be born preterm (<37 weeks' gestation) or low birth weight (LBW; <2,500 g)—outcomes that contribute to racial disparities in morbidity and mortality throughout the life span. The persistent racial gap in birth outcomes has been largely unexplained by individual-level sociodemographic characteristics, psychosocial factors, behavioral risk factors, or biomedical conditions (1–4). A growing body of literature controlling for various neighborhood characteristics (usually census tract or block group level), including socioeconomic status (SES), social support, crime, segregation, pollution, and proximity to supermarkets and health care, has also failed to fully account for racial disparities in birth outcomes (5–15). However, we are aware of only 1 published study that has accounted for racial inequalities in all neighborhood characteristics, both observed and unobserved. In a study of Chicago, Illinois, neighborhoods, Sastry and Hussey (8) explained 30% of black-white differences in continuously measured birth weight with census tract fixed effects but did not partition the proportion attributable to neighborhood context versus individual-level sociodemographic composition or identify important neighborhood features. Moreover, population variation in birth weight may be less meaningful than differences in pathologic markers, such as LBW and its etiologic components, preterm birth (PTB) and growth restriction (being small for gestational age (SGA)), which are more predictive of morbidity and mortality.

Studies of several other outcomes (e.g., hypertension and self-reported health status) have shown that racial disparities can be partially if not fully explained by conditioning on the neighborhood environment in fixed-effects models (16–18). This may be a promising approach for studies of racial disparities, given the inability of available census data (e.g., aggregated SES/deprivation) to wholly characterize relevant neighborhood features (e.g., the built environment, social cohesion, access to goods and services) that are likely to vary by racial composition as a result of residential segregation.

Our objective in this analysis was to determine the total contribution of neighborhood inequalities, both observed and unobserved, to the black-white gap in birth outcomes (LBW, PTB, and term SGA) using a hybrid fixed-effects approach (19, 20) to compare only black women and white women who lived in the same neighborhoods. By comparing these results with those obtained from the conventional approach of controlling only for neighborhood SES, we also evaluated the contribution of unobserved neighborhood factors in explaining racial disparities (i.e., the added value of a hybrid fixed-effects approach).

MATERIALS AND METHODS

Study population

Data for this analysis came from 1999–2001 birth certificate records for Durham and Wake counties in North Carolina, which were used in a larger 8-site study of contextual factors and PTB (21). Singleton non-Hispanic births to black and white mothers were selected for analysis. Women with nongeocodable addresses (3.0%), missing or implausible birth weight or gestational age values (birth weight <500 g; gestational age <20 weeks or >44 weeks), implausible combinations of birth weight and gestational age (0.9%) (22), or missing data on sociodemographic characteristics (0.9%) were excluded from the analysis, for an overall inclusion rate of 96.3% (n = 31,489; 21,221 white women and 10,268 black women).

Outcomes and covariates

Outcomes examined included LBW (<2,500 g), PTB (<37 weeks' gestation), and term SGA (<10th percentile of birth weight for gestational age at ≥37 weeks) (23), all following standard definitions. Given the probable underestimation of SGA at preterm gestational ages (24) and the known association between race and PTB, we avoided differential misclassification by analyzing SGA among term births (37–44 weeks) only. US Census block group (n = 390) was used to define "neighborhood" of residence. Block groups are delineated by local officials and contain an average of 1,500 residents (25). Although socioeconomic indicators at both the census tract and block group levels have shown equivalent associations with LBW (26, 27), the smaller block group was chosen as the neighborhood unit, since there is less potential for environmental heterogeneity within a smaller unit. A validated neighborhood deprivation index, composed of multiple sociodemographic domains, was used as an indicator of neighborhood SES (28). The index was constructed from a principal-component analysis of 8 variables from the 2000 Census of Population and Housing, aggregated by block group, that represented the domains of income/poverty, education, employment, occupation, and housing (e.g., percentage in poverty, percentage unemployed). Individual-level maternal control factors included age (<20, 20–24, 25–34, or ≥35 years), years of education (<12, 12, 13–15, or ≥16), marital status (married or unmarried), and gravidity (first pregnancy, second or third pregnancy, or fourth pregnancy or higher). Maternal behaviors such as prenatal care utilization and smoking were not examined, given that they are less accurately determined from the birth certificate (29, 30) and may be mediators of neighborhood effects (31–33).

Analytic methods

For each outcome, we constructed 4 logistic models to accomplish the following: 1) determine the total contribution of neighborhood to the racial disparity and 2) compare the performance of a neighborhood fixed-effects estimator (controlling for all observed and unobserved neighborhood factors) with that of simple control for neighborhood SES (a conventional practice that controls only for observed neighborhood SES).

Model 1 is unadjusted:

$$\text{logit}(Y_i) = \beta_0 + \beta_1 \text{RACE}_i, \quad (1)$$

where RACE is a binary variable coded 1 for black women and 0 for white women.

Model 2 adjusts for individual characteristics:

$$\text{logit}(Y_i) = \beta_0 + \beta_1 \text{RACE}_i + \beta_2 X_{2i} \ldots \beta_k X_{ki}, \quad (2)$$

where X_{2i} to X_{ki} are $k - 1$ individual-level confounders, such as maternal age.

A fixed-effects or hybrid fixed-effects approach (obtained in a random-effects model) controls for all neighborhood-level variation in both the outcome and exposure(s) (19, 20, 34–36). Model 3 is a hybrid fixed-effects model that adds control for all neighborhood-level confounding of race by adjusting for the neighborhood mean of race in a random-intercept model. The hybrid fixed-effects specification accounts for all between-neighborhood variation in risk that is associated with race and thus provides a within-neighborhood race contrast:

$$\begin{aligned}\text{logit}(Y_{ij}) = {} & \beta_{00} + \beta_1 \text{RACE}_{ij} + \gamma_1 \overline{\text{RACE}}_j \\ & + \beta_2 X_{2ij} \ldots \beta_k X_{kij} + \mu_j, \mu_j \sim N(0, \tau^2), \quad (3)\end{aligned}$$

where $\overline{\text{RACE}}_j$ is the neighborhood mean of race in the study population (i.e., percentage of births to black mothers) and μ_j is the normally distributed random effect specific to each of the j neighborhoods, which has mean 0 and variance τ^2. This model is fitted by the maximum likelihood method, using Gauss-Hermite quadrature to estimate the likelihood function (37). The hybrid model separates the individual-level or within-neighborhood effect of race (β_1) from the contextual or between-neighborhood effect of racial segregation

(γ_1), which can be interpreted as the excess log odds of an adverse birth outcome associated with living in a neighborhood with a higher proportion of black births, regardless of whether an individual is black or white. We compared the race coefficient between hybrid fixed-effects models (model 3) and conventional fixed-effects models (20, 38) obtained from conditional logistic regression with a Hausman test t statistic (20, 39, 40), since other model covariates that are individual-level confounders of race might also be confounded by neighborhood and require separate cluster-mean adjustment. For model parsimony, we did not automatically incorporate a cluster mean for every variable. If the Hausman test was rejected, indicating inconsistency between the fixed-effects and hybrid models, we added cluster-mean adjustment for other covariates that were inconsistent between models and confounding the race effect.

$$\text{Hausman } t \text{ statistic} = \frac{\beta_{FE} - \beta_{RE}}{\sqrt{SE(\beta_{FE})^2 - SE(\beta_{RE})^2}},$$

where FE represents fixed effects, RE represents random effects, and SE represents the standard error. We preferred the hybrid fixed-effects model to the conventional fixed-effects model because it is typically more efficient, does not exclude neighborhoods with only 1 observation or nonvarying outcomes, and can incorporate the effects of neighborhood-level factors (19, 20). Thus, we used the random-effects model and consistent samples to evaluate both study objectives: 1) the total neighborhood contribution to racial disparities and 2) the contribution attributable to neighborhood SES.

Model 4 controls only for neighborhood SES (deprivation index) beyond individual-based adjustment in a random-intercept model:

$$\begin{aligned}\text{logit}(Y_{ij}) = {} & \beta_{00} + \beta_1 \text{RACE}_{ij} + \gamma_1 \text{SES}_{1j} + \gamma_2 \text{SES}_{2j} \\ & + \gamma_3 \text{SES}_{3j} + \beta_2 X_{2ij} \ldots \beta_k X_{kij} \\ & + \mu_j, \mu_j \sim N(0, \tau^2), \end{aligned} \qquad (4)$$

where SES_j refers to quartile of the neighborhood deprivation index.

Transformations to absolute risk differences and relative risks were conducted for all logistic models based on the average covariate values observed in the total population (41). This constitutes a marginal effect at the mean and provides disparity estimates that would occur if both the black and white populations had the covariate distribution of the total population. To determine the neighborhood contribution to disparities, beyond individual sociodemographic characteristics, we examined the incremental reduction in the racial disparity by comparing the difference between models 1 and 3 (model 3 – model 1) with the difference between models 1 and 2 (model 2 – model 1). Although individual sociodemographic characteristics, such as age at childbearing or maternal educational attainment, may be influenced by neighborhood factors, we could not establish temporality, and it is more conservative to assume that sociodemographic factors influence individual selection into neighborhoods and should be controlled as confounders. In a longitudinal analysis of neighborhood effects on self-reported health, adjustment for contemporaneous sociodemographic factors in this fashion was shown to offer a middle ground between adjusting for potential mediators (lower bound) and adjusting for baseline characteristics only (upper bound) (42).

To determine the added value of a hybrid fixed-effects approach versus control for neighborhood SES in a random-intercept model, we compared the race coefficients in models 3 and 4 using a Hausman test t statistic (20, 39, 40). A significant Hausman test statistic indicates the presence of bias in random-effects estimators due to confounding of the exposure variable with the random intercept (a violation of the model assumption). All analyses were performed using Stata 10 software (43).

RESULTS

Clear racial differences in sociodemographic characteristics and neighborhood deprivation were noted (Table 1). Black women tended to be younger, less educated, unmarried, and of higher gravidity than white women. Almost half (48%) of black women lived in neighborhoods from the highest quartile of deprivation, as compared with only 7% of white women. Black women also lived in neighborhoods where the majority of births were black, on average (59%), compared with an average of 20% black births in the neighborhoods of white women—indicative of segregation or racial imbalance in neighborhoods. However, nearly 85% of the study population lived in neighborhoods with at least 5 births to women of each race, which reflects the ability to compute within-neighborhood racial contrasts (not shown).

In unadjusted models (Table 2, model 1), black women were at least twice as likely to deliver LBW, preterm, and term SGA infants compared with white women. On an absolute scale, the disparity was highest for term SGA (8.6%), followed by LBW (7.4%) and PTB (6.2%). On a relative scale, the disparity was highest for LBW (relative risk (RR = 2.8), followed by term SGA (RR = 2.5) and PTB (RR = 1.9). Adjustment for sociodemographic characteristics (Table 2, model 2) reduced these disparities by 31%–43%, depending on the outcome and scale. Reductions were greater for PTB than for LBW or term SGA. Controlling for all neighborhood factors in hybrid fixed-effects models that compared only black and white women who lived in the same neighborhoods (Table 2, model 3) reduced the unadjusted disparities by an additional 4%–15% beyond that explained by individual-level characteristics. This neighborhood contribution to disparities was greatest for PTB (15%), somewhat modest for LBW (7%–8%), and least for term SGA (4%–5%).

Control for only neighborhood SES in random-intercept models (Table 2, model 4) provided disparity reductions that were 2%–6% lower than the hybrid fixed-effects approach, with the largest difference being observed for PTB. Neighborhood SES accounted for only 56% of the total neighborhood contribution to the PTB disparity. The Hausman test for comparison between the hybrid fixed-effects model and the random-effects estimator with neighborhood SES

Table 1. Maternal and Neighborhood Characteristics (%) of Participants in a Study of Neighborhood Inequality and Perinatal Outcomes, by Race, Durham and Wake Counties, North Carolina, 1999–2001[a]

	Black Women (*n* = 10,268)	White Women (*n* = 21,221)
Maternal age, years		
<20	14.4	3.0
20–24	29.4	10.5
25–34	45.3	65.1
≥35	10.9	21.3
Education, years		
<12	20.7	4.8
12	30.4	13.1
13–15	25.2	19.9
≥16	23.7	62.2
Unmarried	58.6	9.5
Gravidity		
First pregnancy	29.4	37.9
Second or third pregnancy	47.4	49.6
Fourth pregnancy or higher	23.3	12.6
Quartile of neighborhood deprivation index[b]		
1 (lowest)	9.1	42.3
2	18.4	31.8
3	24.6	19.0
4 (highest)	48.0	7.0
Average % of black births	59.4 (30.3)[c]	19.7 (18.3)

[a] All associations between characteristics and race were statistically significant at $P < 0.001$.
[b] The neighborhood deprivation index was constructed from a principal-component analysis of 8 variables from the 2000 Census of Population and Housing, aggregated by block group, representing the domains of income/poverty, education, employment, occupation, and housing (28).
[c] Numbers in parentheses, standard deviation.

control rejected the null hypothesis of equivalence of the race effect for PTB only, suggesting a significant unobserved neighborhood contribution to disparities for PTB.

To further explore the neighborhood contribution to racial disparities in PTB, we examined effects by degree of prematurity: moderately preterm birth (MPTB; 32–36 weeks' gestation) and very preterm birth (VPTB; <32 weeks' gestation) (Table 3). These outcomes were modeled according to infants at risk, such that MPTB was contrasted only with term birth (since infants with VPTB are no longer at risk of MPTB), while VPTB was contrasted with both MPTB and term birth. While the absolute black-white risk difference for MPTB (4.2%) was twice that of the much less common VPTB (2.2%), the relative risk was much greater for the rarer, more severe outcome of VPTB (RR = 4.2) than for MPTB (RR = 1.7) (Table 3, model 1). For MPTB, neighborhood accounted for an additional 19%–20% reduction in the racial disparity beyond individual sociodemographic characteristics, yielding an overall disparity reduction of more than 60% (individual + neighborhood; model 3 vs. model 1). For VPTB, disparities were reduced an additional 7%–10% beyond individual characteristics with a neighborhood hybrid fixed-effects model, for an overall disparity reduction of 42%–44% (model 3 vs. model 1). Thus, both individual and neighborhood factors explained less of VPTB disparities than of MPTB disparities. Neighborhood SES could completely account for the smaller neighborhood contribution to VPTB (model 4 vs. model 3). For MPTB, by contrast, the adjustment for neighborhood SES was inadequate to account for all neighborhood factors as in a fixed-effects specification (model 4 vs. model 3). The hybrid fixed-effects model accounted for an additional 10% beyond control for neighborhood SES, which the Hausman test deemed a significant level of bias reduction. In further examinations, other single-variable neighborhood SES measures (i.e., median household income or poverty) performed somewhat less favorably than the composite deprivation index, yielding disparity reductions for PTB that were 5% lower than the deprivation index (not shown).

DISCUSSION

We believe that this is the first study to evaluate the total contribution of neighborhood factors to racial disparities in multiple perinatal outcomes. The results revealed meaningful differences in the neighborhood contribution to racial disparities across perinatal outcomes. For term SGA, there was virtually no neighborhood contribution beyond individual-level sociodemographic characteristics. For LBW, there was a marginal contribution of neighborhood factors that could be entirely explained by a neighborhood SES indicator. Only for PTB was there a notable neighborhood component to racial disparities (15%), beyond the contribution of individual sociodemographic characteristics (40%–43%). The neighborhood component was more strongly observed at moderate ranges of PTB (20% at 32–36 weeks) than for VPTB (7%–10% at <32 weeks). A comprehensive neighborhood socioeconomic index explained only half of the total neighborhood contribution to racial disparities in MPTB, suggesting an impact of racial inequalities in neighborhood environments that is not just a function of measured socioeconomic disadvantage. Several studies have shown greater levels of crime and environmental toxins and lower levels of access to quality health care and supermarkets in racially segregated neighborhoods, even after accounting for neighborhood socioeconomic disadvantage (44–47). Additional study will be necessary to determine the actual integral neighborhood variables, as opposed to aggregate SES (e.g., crime, pollution, food quality), that contribute to black-white perinatal disparities and could be modified by interventions. Alternatively, if neighborhood inequalities related to race did not exist, there would be no neighborhood contribution to disparities. Thus, addressing racial segregation could reduce disparities without identifying the precise neighborhood factors and mechanisms at play.

The greater contribution of individual and neighborhood factors to PTB disparities may reflect stronger social or environmental components of early delivery relative to

Table 2. Unadjusted and Adjusted Black-White Disparities in Perinatal Outcomes, Durham and Wake Counties, North Carolina, 1999–2001

Birth Outcome and Race	Model 1 (Unadjusted)			Model 2[a] (Adjusted)					Model 3[b] (Adjusted Neighborhood Hybrid Fixed Effects)					Model 4[c] (Adjusted Random Effects With Control for Neighborhood SES)				
	%	RR	95% CI	%	RR	95% CI	% Change[d]	95% CI	%	RR	95% CI	% Change[d]	95% CI	%	RR	95% CI	% Change[d]	95% CI
Low birth weight ($n = 31,489$)																		
Black	11.5			9.3					8.9					9.0				
White	4.1			4.3					4.4					4.3				
Risk difference	7.4		6.7, 8.1	5.0		4.3, 5.8	−32	−39, −25	4.6		3.7, 5.4	−39	−47, −30	4.7		3.9, 5.5	−37	−44, −29
Relative risk		2.8	2.6, 3.1		2.2	2.0, 2.4	−35	−42, −28		2.0	1.8, 2.3	−43	−52, −33		2.1	1.9, 2.3	−41	−49, −32
Preterm birth ($n = 31,489$)																		
Black	13.1			11.0					10.3					10.5				
White	6.9			7.3					7.5					7.4				
Risk difference	6.2		5.4, 6.9	3.7		2.9, 4.5	−40	−49, −32	2.8		1.9, 3.7	−55	−67, −43	3.1		2.3, 4.0	−49	−60, −38
Relative risk		1.9	1.8, 2.0		1.5	1.4, 1.6	−43	−52, −34		1.4	1.2, 1.5	−58	−71, −46		1.4	1.3, 1.6	−53	−63, −42
Small for gestational age at term ($n = 28,681$)																		
Black	14.2			11.8					11.5					11.7				
White	5.6			5.8					5.9					5.8				
Risk difference	8.6		7.8, 9.4	6.0		5.1, 6.9	−31	−37, −24	5.6		4.6, 6.6	−35	−43, −26	5.9		5.0, 6.9	−31	−39, −23
Relative risk		2.5	2.3, 2.7		2.0	1.8, 2.2	−33	−41, −25		2.0	1.8, 2.2	−38	−47, −28		2.0	1.8, 2.2	−34	−44, −24

Abbreviations: CI, confidence interval; RR, relative risk; SES, socioeconomic status.

[a] Adjusted for maternal age, education, marital status, and gravidity.

[b] Random-intercept model with adjustment for maternal age, education, marital status, gravidity, and neighborhood racial composition.

[c] Random-intercept model with adjustment for maternal age, education, marital status, gravidity, and neighborhood deprivation index.

[d] Percent change from unadjusted model (model 1); bootstrap confidence interval from 1,000 iterations.

Table 3. Unadjusted and Adjusted Black-White Disparities in Preterm Birth, Wake and Durham Counties, North Carolina, 1999–2001

Birth Outcome and Race	Model 1 (Unadjusted)			Model 2[a] (Adjusted)					Model 3[b] (Adjusted Neighborhood Hybrid Fixed Effects)					Model 4[c] (Adjusted Random Effects With Control for Neighborhood SES)				
	%	RR	95% CI	%	RR	95% CI	% Change[d]	95% CI	%	RR	95% CI	% Change[d]	95% CI	%	RR	95% CI	% Change[d]	95% CI
Moderately preterm birth (32–36 weeks) ($n = 31{,}041$)																		
Black	10.5			8.8					8.2					8.5				
White	6.3			6.5					6.8					6.7				
Risk difference	4.2		3.5, 4.9	2.3		1.5, 3.0	−46	−58, −34	1.5		0.6, 2.3	−65	−82, −49	1.9		1.1, 2.7	−55	−70, −41
Relative risk		1.7	1.5, 1.8		1.3	1.2, 1.5	−48	−60, −36		1.2	1.1, 1.3	−68	−84, −51		1.3	1.1, 1.4	−58	−73, −43
Very preterm birth (<32 weeks) ($n = 31{,}489$)																		
Black	2.9			2.3					2.0[e]					2.0				
White	0.7			0.7					0.7					0.7				
Risk difference	2.2		1.9, 2.6	1.5		1.1, 1.9	−32	−44, −20	1.3		0.9, 1.7	−42	−53, −32	1.3		0.9, 1.7	−43	−53, −33
Relative risk		4.2	3.4, 5.1		3.0	2.3, 3.8	−37	−50, −24		2.8	2.1, 3.5	−44	−61, −28		2.8	2.1, 3.5	−45	−60, −30

Abbreviations: CI, confidence interval; RR, relative risk; SES, socioeconomic status.

[a] Adjusted for maternal age, education, marital status, and gravidity.

[b] Random-intercept model with adjustment for maternal age, education, marital status, gravidity, and neighborhood racial composition.

[c] Random-intercept model with adjustment for maternal age, education, marital status, gravidity, and neighborhood deprivation index.

[d] Percent change from unadjusted model (model 1); bootstrap confidence interval from 1,000 iterations.

[e] Random-intercept model with adjustment for maternal age, education, marital status, gravidity, neighborhood gravidity, and neighborhood racial composition.

SGA or growth restriction. These findings are consistent with those of 2 other studies at the individual level that showed a stronger socioeconomic component to black-white disparities in PTB than in SGA (48, 49). A broader 8-site study, of which these data are part, also showed greater effects of neighborhood deprivation on PTB than on SGA (21, 50). This finding could indicate a stronger social stress mechanism for PTB than for SGA, which has a larger behavioral component via nutrition and substance use. However, the greater contribution of individual socio-demographic and neighborhood factors to disparities in MPTB versus VPTB might also reflect greater homogeneity in obstetric care and practice, according to neighborhood or sociodemographic characteristics, given the greater component of medical intervention for delivery at late preterm gestational ages (51–53).

In the only other study to employ a neighborhood fixed-effects approach to the investigation of perinatal disparities in birth outcomes, Sastry and Hussey (8) accounted for approximately 50% of black-white disparities in birth weight in Chicago with a model similar to ours (model 3) that controlled only for maternal sociodemographic characteristics and neighborhood fixed effects. Although the authors did not assess the contribution of neighborhood fixed effects above and beyond individual sociodemographic characteristics, their total percentage of racial disparities explained was higher than what we observed for LBW in North Carolina (39%–43%). For comparability, we repeated the analysis accounting only for neighborhood fixed effects on continuous birth weight and still found a lower percentage explained in North Carolina (15% in North Carolina vs. 30% in Chicago), even though the crude disparity was similar in both locations (approximately 325 g). It is important to recognize that the neighborhood component of disparities is directly proportional to the degree of racial inequality in neighborhoods; thus, areas with greater segregation, like Chicago, may have a larger neighborhood contribution than we observed in North Carolina. In the 2000 US Census, the isolation segregation index in the Chicago Metropolitan Statistical Area was 0.78 (54) as compared with 0.57 and 0.38 in Durham and Wake counties (55), respectively. Nonetheless, Sastry and Hussey's findings for birth weight (8) are broadly similar to our results for PTB.

By controlling for the racial composition of neighborhood of residence, other investigators may have inadvertently obtained a neighborhood fixed effect of racial disparities in birth outcomes through a hybrid specification in random-effects models (10, 14). Additionally, in Chicago, Buka et al. (10) found that the black-white gap in birth weight dropped an additional 22% when the authors additionally controlled for proportion black and a disadvantage index, beyond maternal sociodemographic and behavioral factors, with a total reduction of 66%. These figures are comparable to what we observed in our study for MPTB. While only the disadvantage index was predictive when results were controlled for proportion black, neighborhood disadvantage may mediate the effect of racial residential segregation, and it can be inappropriate to model collinear variables (56). Another analysis carried out in New York City similarly found that a tract-level segregation measure explained 22% of the black-white LBW disparity in models adjusted for sociodemographic and behavioral factors (14). While neighborhood poverty did not help to explain the effect of segregation in the New York City study, there was a significant cross-level interaction wherein racial disparities were not observed in high-poverty tracts. In North Carolina, we examined potential cross-level interactions between race and neighborhood deprivation or segregation (percentage of black births) but did not find substantive neighborhood variability in the effect of race or meaningful interactions with either neighborhood variable.

Overall, we found that control for individual sociodemographic factors and all neighborhood factors yielded total reductions in the black-white disparity of approximately 40% for LBW, term SGA, and VPTB, 60% for PTB, and 70% for MPTB, which leaves one-third or more of the crude disparities unexplained. Unfortunately, less of the racial disparity in VPTB (40%) than the disparity in MPTB (70%) could be explained in our analysis, and VPTB accounts for the largest component of disparities in infant mortality (57). Given that the observed neighborhood deprivation index was actually more strongly related to VPTB than to MPTB, it appears that the determinants of the racial disparity in VPTB are different from those of MPTB and are less well captured by the sociodemographic and neighborhood factors measured in our study. In further sensitivity analyses, additional control for the available behavioral factors, prenatal care utilization and smoking, did not reduce observed racial disparities or the neighborhood contribution to disparities.

This study had several strengths in evaluating the total contribution of contemporaneous neighborhood inequalities to racial disparities in multiple perinatal outcomes. We reported both absolute and relative disparities and used population-based data rather than a clinic-based sample that could have been nonrepresentative. The analysis of metropolitan counties in a southern state complements the existing literature on northern cities, and the block-group definition of neighborhood also provided finer control for the immediate residential environment. By employing random-effects models to obtain fixed-effects results in the hybrid specification, we were able to include all observations and to evaluate between-neighborhood effects of racial segregation and deprivation (not shown), avoiding some of the limitations of fixed-effects models. This is a simple technique that can be easily implemented by researchers analyzing clustered data in space or time who want to estimate both within-cluster effects and between-cluster effects (19, 20, 34). It is important to recognize that when multiple level-1 confounders are included in this approach, they may also require cluster-mean adjustment to obtain the actual within-cluster effect of the variables of interest (20). In our study, we were generally able to obtain estimates equivalent to those of a conventional fixed-effects specification by adding only the neighborhood mean of race. However, for the outcome of VPTB, we also had to add the neighborhood mean of gravidity (percentage of each category), since it was confounded by neighborhood and also confounded the effect of race. This could be a limitation of other studies that do not compare their random-effects results with a fixed-effects specification. Differences in the coefficients between random- and fixed-effects models indicate a violation of a random-effects model assumption—that the random neighborhood intercepts are uncorrelated with exposures (37–40). This leads to neighborhood confounding of exposures in random-effects models.

Although our approach provides complete control for neighborhood inequalities resulting from residential segregation, by examining disparities within neighborhoods, our estimate of the neighborhood component of disparities is sensitive to the control of all individual-level confounders of neighborhood selection and only accounts for contemporaneous neighborhood exposure. While we controlled for several potential sociodemographic determinants of neighborhood choice, data on some factors were not available on the birth certificate (e.g., income, wealth), and some of the factors we did control for (e.g., education, age at childbearing) may have been influenced by prior neighborhood conditions. Thus, our results may have been at risk of both underadjustment for confounders (upward bias) and overadjustment for mediators (downward bias). The methods that account for simultaneous mediators and confounders can only be used with longitudinal data. Nonetheless, our cross-sectional adjustment strategy may yield estimates that fall between adjustment for potential mediators (lower bound) and adjustment for prior SES only (upper bound) (42).

Moreover, we were unable to consider time-varying or life-course dimensions for individual or neighborhood factors, which may be critical to understanding disparities. SES in both childhood and adulthood has been shown to influence the likelihood of delivering an LBW infant (58–60), and blacks are more likely to experience a longer duration of poverty at both the individual and neighborhood levels (61–64). This means that black women currently living in high-SES neighborhoods are more likely to have spent time living in low-SES neighborhoods and that black women currently living in low-SES neighborhoods are more likely to have lived there longer than whites. Thus, cross-sectional data may underestimate socioeconomic differentials, leading to residual confounding of the race effect. In a recent analysis, cumulative measures of individual and neighborhood poverty were found to explain a substantially greater proportion of disparities in self-rated health than single time-point measures (61). Only 1 data set in Illinois has been used to examine the contribution of SES at more than 1 time point to black-white differences in birth outcomes. Disparities in LBW were found to persist among women who lived in high-SES neighborhoods at birth and delivery, but individual-level income was not controlled, and the definition of high SES ranged from \$30,000 to \$150,000 in neighborhood median family income, raising the concern of residual confounding due to residential segregation (11, 65). Evaluating life-course differentials in individual and neighborhood factors may be a promising approach to the study of racial disparities in birth outcomes.

In summary, we find that contemporaneous neighborhood inequalities account for a notable portion of black-white perinatal disparities only in PTB (approximately 15%), particularly in MPTB (approximately 20%), in 2 counties of North Carolina. A comprehensive neighborhood socioeconomic index could not completely account for the neighborhood contribution to PTB disparities. Therefore, the fixed-effects approach appears to add value, and researchers should be careful to test the full contribution of neighborhoods against the specific neighborhood variable(s) examined. Additional research is needed to determine the integral neighborhood factors that could be modified in interventions to address disparities. Alternatively, approximately 15% of the PTB disparity might be averted by eliminating racial segregation. Nonetheless, one- to two-thirds of black-white gaps in birth outcomes remain unexplained by contemporaneously measured individual sociodemographic and neighborhood factors. In future studies, researchers should evaluate other differential exposures (e.g., racism or wealth) and neighborhood factors across the life course.

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