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Neighborhood ethnic density and preterm birth across seven ethnic groups in New York City

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Abstract

Residential segregation limits non-white ethnic groups' access to white neighborhood resources, but may also reduce their exposure to discrimination and facilitate social support. We computed adjusted preterm birth risk differences (RDs) for seven ethnic groups comparing > 25% to ≤ 25% ethnic density neighborhoods using 1995–2003 New York City birth records and a spatial ethnic density measure. RDs ranged from −15.0 per 1000 (95% CI: −18.5, −11.4) for whites to 6.4 per 1000 (95% CI: 2.8, 9.9) for blacks, with Hispanic and Asian estimates falling in between but tending to be protective. Results suggest that ethnic density is uniquely harmful for non-Hispanic blacks.

Keywords

Race; Ethnicity; Ethnic density; Residential segregation; Preterm birth; Birth outcomes

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Appendix A. Supplementary material: Supplementary data associated with this article can be found in the online version at doi: 10.1016/j.healthplace.2010.11.006.

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1. Background

Racial and ethnic residential segregation is deeply entrenched and widespread in the social geography of the United States (Massey and Denton, 1987; Fong and Shibuya, 2005; Massey and Denton, 1989), with some areas so segregated that they have been compared to apartheid-era South Africa (Massey, 1990). Social and health science research in the US has documented links between segregation and a variety of social and physical ills in the black population (Cutler and Glaeser, 1997; Grady, 2006; Morenoff, 2003; Ellen, 2000; Baker and Hellerstedt, 2006; Polednak, 1991; Laveist, 1989; Polednak, 1996; Guest et al., 1998; Collins and Williams, 1999; Cooper et al., 2001; Hart et al., 1998; Jackson et al., 2000; Huie et al., 2002; LeClere et al., 1997; Polednak, 1993; McCord and Freeman, 1990; White and Borrell, 2006; Subramanian et al., 2005; Morello-Frosch and Jesdale, 2006; Acevedo-Garcia, 2001; Chang, 2006; Peterson and Krivo, 1999; Mason et al., 2009). From a perspective that privileges material resources as the means to health (Lynch, 2000), the observed harms of segregation are unsurprising, given the historical discrimination of blacks in employment and education that segregation has facilitated. From a psychosocial standpoint (Pickett and Wilkinson, 2008), however, black neighborhoods might benefit their residents by limiting stigmatizing inter-racial interactions, facilitating social networks, or providing a context for political organizing (Sampson and Groves, 1989; Hutchinson et al., 1996; Putnam, 2007; Bledsoe et al., 1995). Indeed, a handful of epidemiologic studies have documented better black birth outcomes (Bell et al., 2006; Roberts, 1997) and lower black mortality (Fang et al., 1998; Inagami et al., 2006) in majority-black compared to more heterogeneous neighborhoods.

Among less-studied groups such as Hispanic and Asian immigrants, for whom segregation may arise more from patterns of chain migration than from historical oppression, ethnic density may be less detrimental than it is among black Americans (Morenoff, 2003; Jenny et al., 2001; Gorman, 1999; Osypuk et al., 2009; Eschbach et al., 2005, 2004; Ostir et al., 2003; Patel et al., 2003; Reyes-Ortiz et al., 2009; Kandula et al., 2009; Walton, 2009). Segregated neighborhoods may also provide unique protections to more recently-arrived ethnic groups by buffering the stress of acculturation and providing access to country-of-origin foods (Lieberson, 1961; Duany, 1998).

The aim of this paper is to increase understanding of the segregation-health relationship by examining preterm birth risk in ethnically dense areas across multiple, often understudied, ethnic groups. Preterm birth, or birth before the 37th week of gestation, is an outcome of particular public health relevance because it is an important cause of infant mortality, leads to a variety of morbidities and learning impairments in children and adults, and is the largest contributor to the two-fold black-white disparity in infant death (Berkowitz and Papiernik, 1993; Fiscella, 2004; Hogue and Hargraves, 1995). Understanding of the etiology of preterm birth remains vague, but mounting evidence suggests that stress may play a particularly important role, either by triggering hormones related to labor initiation or through an inflammatory pathway provoked by immune suppression and infection (Culhane and Elo, 2005; Rowley, 1994; Hogue and Bremner, 2005; Wadhwa et al., 2001a; Wadhwa et al., 2001b; Goldenberg et al., 2008). Several studies have documented a correlation between preterm birth and stressful life experiences (Dole et al., 2004; Lu and Chen, 2004; Collins et al., 2000; Rosenberg et al., 2002), self-reported stress (Rini et al., 1999), and stressful neighborhood environments (Messer et al., 2006a). These studies suggest that preterm birth may be sensitive to the material and psychosocial environments that vary with neighborhood ethnic density.

We used New York City birth records to estimate the risk of preterm birth associated with neighborhood ethnic density. Studies of ethnic density effects for Hispanics and Asians remain rare; to address this gap in the literature, we included Spanish Caribbeans, Central Americans (plus Mexicans), South American Hispanics, East Asians, and South Asians in our analyses, in addition to non-Hispanic whites and blacks. Furthermore, most studies to date have relied on administrative units (e.g. census tracts) alone to define neighborhood ethnic density (so-called “aspatial measures”). Because of concerns that aspatial measures may mis-characterize the geographic distributions of ethnic groups (Reardon, 2006), we represented the ethnic composition of each mother's neighborhood with a spatial measure, “proximity-weighted ethnic density,” which allows the ethnic composition of the areas surrounding the mother's residence to influence her estimated exposure in proportion to their distance from her (Reardon and Firebaugh, 2002; Reardon et al., 2008).

2. Methods

2.1. Data Sources and management

New York City birth records from January 1, 1995 through December 31, 2003 provided outcome (gestational age), ethnicity, and individual-level covariate data on all singleton births occurring to residents of New York City over the nine-year period (*N*=1,052,576). The birth records were geocoded and each observation was assigned a 1990 or 2000 census tract number (depending on the year of birth) by the New York City Department of Health and Mental Hygiene. We excluded births missing gestational age information (*N*=6418). We also excluded births if they were missing census tract or county information (*N*=108,433), or were assigned to non-existent (*N*=1812), ambiguous (*N*=62), or unpopulated (*N*=28) census tracts. Preterm birth rates were similar across observations with and without complete census tract information.

We defined seven ethnic group categories (non-Hispanic white, non-Hispanic black, Spanish Caribbean, Central American plus Mexican, South American Hispanic, East Asian, and South Asian) from the self-reported race and ethnic origin variables available in the birth records (Fig. 1). Births without the race or ethnic origin information necessary to create maternal ethnic group categories (*N*=8801) were excluded. Births to women with non-Hispanic ethnic origins that were not white, black, East Asian, or South Asian (*N*=13,923) or to women who reported a Hispanic ethnic origin not in the Spanish Caribbean, Central America, or South America (*N*=25,212) were not included in the study. These exclusions left 887,887 observations for the analysis.

The NYC Department of Health and Mental Hygiene assigned 1990 tract numbers to births occurring in 1995–1999 and 2000 tract numbers to births occurring in 2000–2003. Between the two censuses, the Census Bureau split some 1990 census tracts to create two 2000 census tracts, and merged some 1990 tract pairs to create single 2000 tracts (New York City Department of City Planning. Table G-1: New York City 1990 Census Tracts with Boundary/Number Changes for 2000 Census, 2001). To create consistent tract numbers over time, we gave 1990 tracts that were merged in 2000 their corresponding 2000 tract numbers. Likewise, we gave 2000 tracts that had split from 1990 tracts their “parent” 1990 tract numbers. After updating, there were 2156 unique tracts containing births included in the analysis.

We obtained total and ethnic group population counts in the 2217 tracts contained in the five counties of New York City from Summary File 1 of the 2000 US Census and area-level covariates from Summary File 3. In order to match the birth records, we merged 15 year–2000 tract pairs to re-create the 15 1990 tracts from which they were split, leaving 2202 unique tracts in the census data. Census tracts that were not found in the birth records consisted primarily of low-population tracts and Tract 1 in the Bronx, corresponding to Riker's Island Prison.

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2.2. Variables and variable construction

We defined the outcome, preterm birth, as a live singleton birth at greater than 20 but less than 37 completed weeks of gestation using the clinical estimate of gestational age (Berkowitz and Papiernik, 1993).

We defined the exposure, neighborhood ethnic density, as the percentage of the population in a woman's area of residence with a given ethnic identity. For non-Hispanic white and black mothers, respectively, we used non-Hispanic white and black densities as the exposures. For Spanish Caribbean, Central American, and South American mothers, we defined the exposure as the neighborhood density of Hispanics, and for East and South Asian mothers we defined the exposure as Asian density. While region-specific ethnic densities (e.g. density of Central Americans) were available from the census data, and were theoretically preferable, they included a large amount of missing data due to small-population data suppression (US Census Bureau. Summary File 4: Technical Documentation, 2000). We did, however, conduct a secondary analysis using the region-specific ethnic densities for comparison.

Following previous authors (Reardon and Firebaugh, 2002; Reardon et al., 2008), we assumed that the areas nearest a woman contributed most to her experience of neighborhood-level ethnic density. We allowed populations farther away to influence her estimated exposure as well, but this influence decreased in proportion to distance. Because they were the smallest unit available in the birth records, we used census tracts to locate the women geographically, and used the distances between approximate census tract centers (centroids) to estimate the distance from each woman's residence to other populations. New York City census tracts are geographically small, with a median area of 0.18 square kilometers, allowing us to locate populations fairly precisely. We positioned centroids using a center-of-mass generator (Jenness Enterprises: Polygon Center of Mass), which estimates the geographically-weighted center of each tract, and computed between-centroid distances in ArcGIS (ESRI).

We calculated the “proximity-weighted ethnic density” (Reardon and Firebaugh, 2002) ($\prod_{JM}$) for a woman belonging to ethnic group *M* and residing in census tract *J* by multiplying the population count of ethnic group *N* in each census tract *K* (x_{KN}) by a weight (p_{JK}) that represents the proximity of tracts *J* and *K*. We summed these weighted ethnic populations and divided the sums by total census tract populations (x_K) that were weighted in the identical manner. This produced a weighted “percent” as shown below:

$$\prod\nolimits_{JM} = \frac{\sum_K (x_{KN} \times p_{JK})}{\sum_K (x_K \times p_{JK})}$$

The proximity weight (p_{JK}), a “biweight kernel”, allows census tract *K*'s influence to decay in an approximately Gaussian manner with its distance from census tract *J* (Lee et al., 2006):

$$p_{JK} = \left(1 - \left(\frac{d_{JK}}{r}\right)^2\right)^2 \quad \text{if } d_{JK} < r, \ \text{else } p_{JK} = 0$$

where d_{JK} is the distance between census tracts *J* and *K*. Note that if *J*=*K*, then d_{JK}=0 and p_{JK}=1; that is, a census tract's own ethnic composition will have maximal influence on the estimated exposure of the residents of that census tract. The variable *r* is the distance from census tract *J* beyond which there is no influence on *J*'s estimated ethnic density. The value of the radius was chosen based on the hypothesized area thought to meaningfully affect the environment of those living in census tract *J*. Lee and colleagues suggest a radius of 500m to approximate residential areas accessible by foot (Lee et al., 2006), which we considered to be an appropriate neighborhood definition for a densely populated urban area such as New York City. Because there is no generally accepted threshold at which ethnic density is thought to be most influential, we dichotomized proximity-weighted ethnic density at 25% in order to allow an adequate number of births in both exposed and unexposed categories across ethnic groups. Two sensitivity analyses were conducted, one with ethnic densities dichotomized at 20% to ensure that results were not driven by random variability at one cut-point, and a second with ethnic densities coded as continuous variables with squared terms allowing for exploration of non-linearities in the ethnic density—preterm birth association.

We included the following covariates in the adjusted models: maternal age (indicators for < 20, 20–34, and 35+ years), education taking age into account (indicators for <12 years and < 20 years of age, < 12 years and 20+ years of age, 12 years, 13–15 years, and 16+ years), nativity (US- or foreign-born), parity (indicators for 1, 2–5, and 6+ previous births), tobacco use (smoker or nonsmoker), prepregnancy weight (indicators for < 125, 125–150, and > 150 pounds), prenatal care received in first 120 days of gestation (yes or no), and payment type (indicators for private insurance, Medicaid, or out-of-pocket).

In addition, we included two neighborhood-level covariates, residential stability (percent of the neighborhood population residing in the same house from 1995 to 2000) and neighborhood deprivation. Neighborhood deprivation was represented using a standardized index arising from 17 tract-level census variables (% of the population with less than a high-school education, % unemployed, % males not in work force, % crowding, % renter-occupied units, % male professionals, % female professionals, % males in management, % females in management, % poverty, % female-headed household with children, % households with < $30,000/year, % households on public assistance, % households with no car, median household income, median income of individuals with earnings, median value of owner-occupied units) that were summarized using principle components analysis as previously described (Messer et al., 2006b). Using this index allowed us to adjust for multiple highly correlated dimensions of neighborhood deprivation in the model without creating problems of multicollinearity. Both residential stability and the component variables of the neighborhood deprivation index were proximity-weighted in the same manner as ethnic density and dichotomized at the overall median. Adjustment for continuous neighborhood deprivation produced similar results.

2.3. Data analysis

We used logistic regression to model the relationship between preterm birth and dichotomized proximity-weighted ethnic density for each ethnic group separately. The Huber-White “sandwich” variance estimator was employed to account for clustering at the census tract level (Williams, 2000). The coefficients from these marginal models closely approximated the results from random-intercept models, for which the estimated intra-cluster correlation coefficients were very small (all < 0.02); we therefore chose to use marginal models. Recent articles have also argued that results from marginal models are more appropriate for public health inference because they estimate an average effect for the entire population rather than for the population of a single neighborhood (Hubbard et al., 2010).

We employed the following modeling strategy for all ethnic groups. First, we modeled the log odds of preterm birth as a function of ethnic density alone to estimate the crude exposure-outcome association. Second, models were adjusted for all covariates. Third, we re-ran adjusted models without the two most frequently missing covariates: prenatal care and prepregnancy weight. Almost 20% of observations were missing one or more of these variables. We conducted a change-in-estimate analysis to assess the extent of confounding incurred by their exclusion; we considered a change in the odds ratio of less than 10% small enough to warrant omitting them to increase precision and generalizability (Mickey and Greenland, 1989). Fourth, we stratified the models by neighborhood deprivation, since the psychosocial correlates of segregation may have a different association with preterm birth depending on the resource environment that is also present (Pickett et al., 2005; Phillips et al., 2009). Finally, we computed crude, adjusted, and stratified risk differences (RDs) from the logistic model regression coefficients, with US-born women aged 20–34 who were high-school educated, had 2–5 previous live births, received early prenatal care, were on Medicaid, and resided in a more stable and poorer neighborhood as the reference risk group. Risk differences provide an estimate of the number of preterm births attributable to (or prevented by) residence in ethnic enclaves (assuming the modeled associations are correct and causal), and are therefore particularly informative for public health and policy applications.

We conducted several sensitivity analyses to assess potential changes in the results when using different population and variable specifications. First, we re-ran the models with ethnic density dichotomized at 20% rather than 25%. Second, we used linked hospital discharge diagnosis and procedure codes to identify and exclude medically indicated preterm births (births subsequent to any surgical or medical induction of labor and births subsequent to pre-labor cesarean sections) in order to obtain results specific to spontaneous preterm birth. Third, we restricted analyses to primiparous women, to remove any influence of repeat births to the same mother over the nine-year study period. Fourth, were-ran models among mothers whose ethnic identity matched the father's ethnic identity, since the father's ethnic affiliation may influence the mother's experience of ethnic density in her neighborhood. Finally, we re-ran models among foreign-born women only.

We also re-ran analyses with region-specific ethnic densities (e.g. Central American density rather than Hispanic density), dichotomized at 15% to accommodate the lower average density of regional populations.

To provide a more complete picture of the shape of the relationship between ethnic density and preterm birth, we plotted predicted probabilities of preterm birth estimated from adjusted models with ethnic density coded continuously with a squared term (to allow for non-linearities). We used these predicted probabilities to calculate risk differences for 30–10% and 50–10% ethnic density contrasts.

3. Results

The majority of the 887,887 births included in the analysis occurred to non-Hispanic white, non-Hispanic black, or Spanish Caribbean women (Tables 1a and 1b), reflecting the ethnic distribution of the city as a whole. The proportion of births to Central Americans, South Americans, and South Asians was greater than their proportion in the general population, indicating high fertility.

East Asians had the lowest risk of preterm birth of all the ethnic groups, followed closely by non-Hispanic whites. Non-Hispanic blacks had by far the highest risk (Tables 1a and 1b). Non-Hispanic blacks did not, however, have the least favorable distribution of covariate risk factors, as they were more likely than Spanish Caribbeans, Central Americans, and South Americans to have education beyond a high school degree, were less likely to be on Medicaid than any other group except whites, and were more likely than Spanish Caribbeans or East Asians to have early prenatal care.

The degree of ethnic density commonly experienced in the maternal neighborhood varied drastically by ethnic group, with non-Hispanic white and black births occurring largely to women residing in majority white or black neighborhoods, respectively (Fig. 2), but East and South Asian births occurring mostly to women in neighborhoods with only a small proportion of other Asians. The Hispanic groups fell in between, with Spanish Caribbean births more likely to occur in highly Hispanic neighborhoods than either Central or South American births. These ethnic density differences reflect the relative size of the ethnic populations, but also follow documented national and historical trends in which blacks and whites are highly segregated from one another, while Asians tend to integrate into white neighborhoods and Hispanics fall somewhere in between (Fischer et al., 2004; Massey and Denton, 1987; Massey, 1981). The geographic distribution of these groups is shown in Fig. 3, which illustrates the high degree of clustering by ethnicity.

Crude changes in preterm birth risk associated with maternal residence in an ethnic neighborhood (> 25% ethnic density) versus a less ethnically dense neighborhood ranged from −17.0 per 1000 (95% CI: −20.9, −13.1) for white women, indicating a substantial protective effect of own-group density, to 9.5 per 1000 (95% CI: 6.0, 13.1) for black women, indicating increased risk associated with residence in a black neighborhood. The Hispanic and Asian group estimates fell between those for whites and blacks. Controlling for covariates moved the estimates toward the null for all groups except South Americans (Table 2, Fig. 4). When adjusted, the risk difference was −15.0 per 1000 (−18.5, −11.4) among whites and 6.4 per 1000 (95% CI: 2.8, 9.9) among blacks.

The two most frequently missing variables—prenatal care and prepregnancy weight—were not included in the final adjusted models, because the change in the odds ratio resulting from their exclusion was 5% or less in all groups. Fully adjusted risk differences (computed with these two variables retained) are presented in Table 1 of the Online Appendix for comparison; estimates from the fully adjusted models were farther from the null for all groups except whites and East Asians. These results should be treated with some caution, however, as they are based on analyses missing over 20% of the observations.

Changes in the risk differences across neighborhood deprivation strata exceeded 5 per 1000 for non-Hispanic white, Central American, South American, and South Asian groups. Interactions between ethnic density and neighborhood deprivation were statistically significant ($p < 0.05$) for non-Hispanic whites and Central Americans. Risk differences for white women in richer and poorer neighborhoods were −8.3 (95% CI: −14.4, −2.2) per 1000 and −20.0 (95% CI: −25.9, −14.1) per 1000, respectively. For Central Americans the risk differences per 1000 were 2.1 (95% CI: −4.2, 8.5) and −9.6 (95% CI: −18.5, −0.8), for South Americans they were 3.2 (95% CI: −3.5, 9.9) and −2.8 (95% CI: −19.3, 13.6), and for South Asians they were −4.9 (95% CI: −11.9, 2.1) and −15.3 (95% CI: −32.0, 1.4) in richer and poorer neighborhoods, respectively. For all the groups, with the exception of non-Hispanic blacks, the RD was lower when estimated in poorer neighborhoods (Table 2, Fig. 5), but many of these estimates were quite imprecise.

We re-ran stratified models for the white, Hispanic, and Asian groups with non-Hispanic black density included, to explore the possibility that differences in estimates across neighborhood deprivation categories are driven by differences in the “out-group” ethnic composition. (For example, white women residing in non-white neighborhoods are more likely to be living with Asians if their neighborhood is wealthy and blacks if their neighborhood is poor.) Controlling for non-Hispanic black density in the models did not, however, change the overall pattern of the results, although some estimates moved slightly toward the null (Online Appendix Table 2).

The overall pattern of findings remained largely unchanged in additional sensitivity analyses (Table 3 and Fig. 1 of the Online Appendix). For the smaller groups (e.g. South Americans), restricting to primiparous women shifted the estimates more substantially, but the level of imprecision was also increased so it was difficult to say whether this was a meaningful change. When the father's ethnic identity matched the mother's, the effect of ethnic density appeared to be less protective among white mothers but more protective among Spanish Caribbean mothers; however, paternal ethnicity information was missing for about 20% of the births, so these results should be interpreted with caution.

When we used region-specific ethnic densities as the exposures, the Central American and South Asian estimates were moved close to the null, but these estimates were obtained from models with 12.5% of Central American observations and 18.4% of South Asian observations excluded due to missing exposure values (two sets of results, one with missing data excluded and the other with missing data assumed to be zero are presented in the Online Appendix Table 4). The null value for the Central American estimate appears to arise at least in part from a scarcity of women in wealthy Central American tracts (N=93), such that it was not possible to adequately control for neighborhood deprivation.

Predicted probabilities estimated using continuous ethnic density, with a squared term to allow for deviations from linearity, suggest that the relationship between ethnic density and preterm birth may be non-linear in some groups (Online Appendix Fig.2 and Table 5). In particular, black preterm birth risks in wealthier neighborhoods appear to increase with ethnic density until ethnic density reaches around 50%, then level off and possibly decrease slightly. Overall, deviations from linearity were small, however, and indicate that the contrast between > 25 % and ≤ 25 % ethnic density from the main analysis are a useful summary of the ethnic density—preterm birth relationship across groups.

4. Discussion

Our study results suggest that higher ethnic density is associated with poorer birth outcomes among non-Hispanic black women, consistent with some (Grady, 2006; Morenoff, 2003; Ellen, 2000) (Baker and Hellerstedt, 2006; Osypuk and Acevedo-Garcia, 2008) but not all (Roberts, 1997) previous research. We found a substantial reduction in preterm birth risk among white women living in whiter neighborhoods, and among most Hispanic and Asian groups, the associations between ethnic density and preterm birth appeared to be null or slightly protective.

A growing body of evidence suggests that ethnic density may positively influence an area's social environment (Pickett and Wilkinson, 2008; Putnam, 2007). Recent experimental research suggests, for example, that the capacity for collective action is influenced by norms of reciprocity that encourage collaboration between members of the same ethnic group and discourage cross-ethnic cooperation (Habyarimana et al., 2009), making ethnic density a potentially important determinant of social organization. Evidence of variation in ethnic density effects that we observe here suggests that hypothesized benefits of ethnic density may accrue more to some ethnic groups than others, however. For example, the difference in ethnic density responses between non-Hispanic blacks and Hispanics, despite similar levels of poverty, may result from unique protections in Hispanic neighborhoods (such as healthful food environments) that are not available to non-Hispanic blacks.

Alternatively, unmeasured differences in neighborhood resources may influence ethnic density effects. In particular, a long history of racial oppression and chronic under-investment in black neighborhoods may make black density a marker of entrenched and concentrated neighborhood poverty that our neighborhood deprivation measures do not fully capture. The historical context in which neighborhoods have often been formed could, in addition, influence the psychosocial well-being of their residents; years of racially-based barriers to geographic mobility may, for instance, create a sense of powerlessness among residents of black neighborhoods. It is also possible that the associations between ethnic density and preterm birth differ because non-Hispanic black women in New York City are more likely to be US-born—and may thus have spent more time in their neighborhoods—than the largely immigrant Hispanic and Asian groups. Although we adjusted for maternal immigrant status, we had no information on timing of neighborhood residence, so we could not account for differences in exposure duration.

Stratifying by neighborhood deprivation suggested that the influence of ethnic density on health may be modified by the material environment. When estimated within poorer neighborhoods, risk differences for all groups (with the notable exception of non-Hispanic blacks) indicated a protective effect of residence in an ethnic neighborhood, although some point estimates were close to null. Among whites and South Asians, particularly, the risk reductions in poorer neighborhoods were sizeable and differed substantially from estimates in wealthier areas. Scarcity of health-promoting resources in poorer neighborhoods may increase the relative importance of psychosocial benefits arising from a shared ethnic or cultural identity. This possibility could not be examined with the available data, however.

Census data suppression for small population groups limited our ability to estimate region-specific ethnic density effects. Region-specific ethnic densities may be more meaningful for the social experience of a woman in a given neighborhood than densities based on broader ethnic definitions. While we provide supplemental estimates of region-specific ethnic density effects, the imprecision of the results and the high level of exposure missingness makes these estimates difficult to interpret. The spatial patterning of births suggests that region-specific groups live in distinct areas of the city, and thus broader ethnic densities may closely approximate the region-specific ethnic density that is predominant in a given neighborhood; for example, because of self-segregation by region, Hispanic density may be a reasonable proxy for Central American density when estimated in a Central American woman's neighborhood. Nonetheless, robust measures of region-or even country-specific densities might reveal additional variation of effects, and future research in this area might consider examining groups, such as Puerto Ricans and Chinese, with potentially sufficient numbers to support this level of nuance.

Recent publications have highlighted the problem of investigating the independent effects of neighborhood economic and ethnic segregation (Oakes, 2006; Messer et al., 2010), since these two characteristics tend to be highly correlated (Sampson et al., 2008). An examination of the underlying distribution of the exposure and covariates within neighborhood deprivation strata revealed few cells with a glaring lack of data. Some uncontrolled confounding is, of course, still possible due to heterogeneity within covariate categories. The difficulty of disentangling the independent influences on health of the neighborhoods themselves from the characteristics that cause individuals to select into the neighborhoods in the first place also remains a challenge (Oakes, 2004).

Despite its limitations, our study has two key strengths. First, we included a large number of ethnicities, including under-studied Hispanic and Asian groups, enabled by the unique diversity of the New York City population. In addition, we employed a spatial measure of neighborhood-level ethnic density to address the documented limitations of non-spatial measures. The radius, 500 meters, represents a walkable distance around the residential area (Lee et al., 2006), and was chosen as a theoretically appropriate neighborhood approximation for a population-dense urban area like New York City.

The results of this analysis suggest that the balance of beneficial and harmful material and psychosocial correlates of segregation may differ across ethnic groups. Segregation appeared to benefit whites and harm blacks in this study, perhaps reflecting the long history of unequal resource distribution between blacks and whites of which segregation is a cause, consequence, and marker. The more recently-arrived groups that are largely outside this history had somewhat more limited responses, but results suggested a protective effect, especially in poorer neighborhoods. The data used for this analysis prevented investigation of hypothesized pathways between ethnic density and health, but suggest that historical context may be important for understanding the associations between neighborhoods and health; the findings provide a basis future research exploring these mechanisms in greater depth.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

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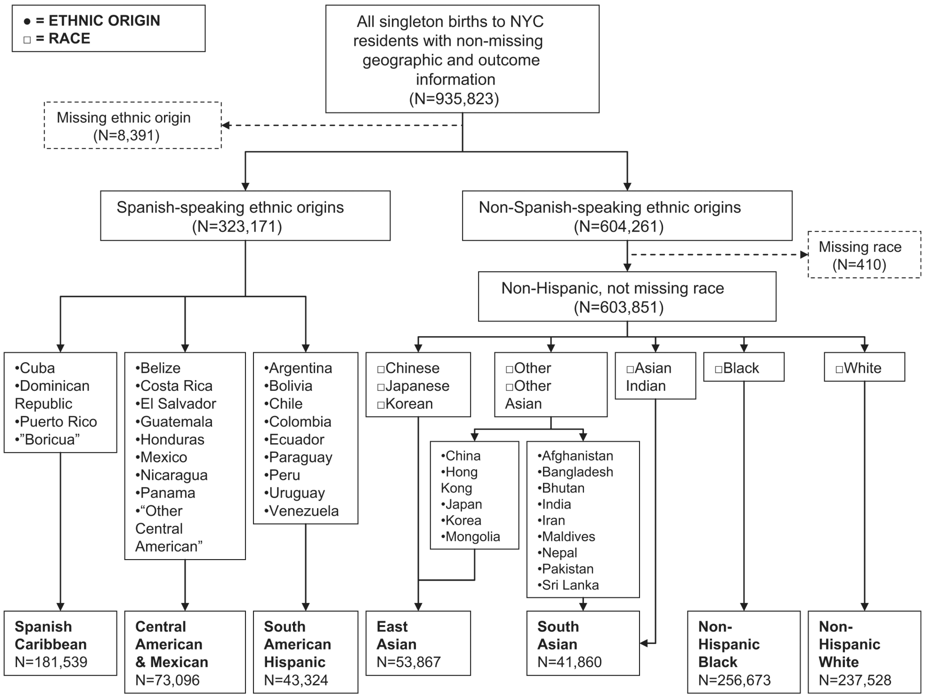

Fig. 1.

Identification of seven maternal ethnic groups from the New York City birth records for 1995 through 2003. Dashed lines indicate births that could not be included due to missing race or ethnicity information.

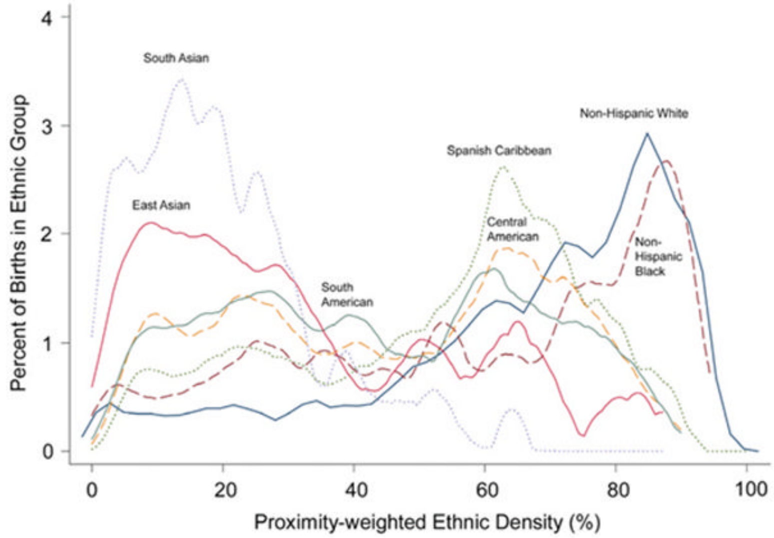

Fig. 2.

Distribution of births in seven ethnic groups across the range of ethnic density in the maternal neighborhood: New York City, 1995–2003. (Kernel smoothed; kernel=Epanechnikov, bandwidth=0.02).

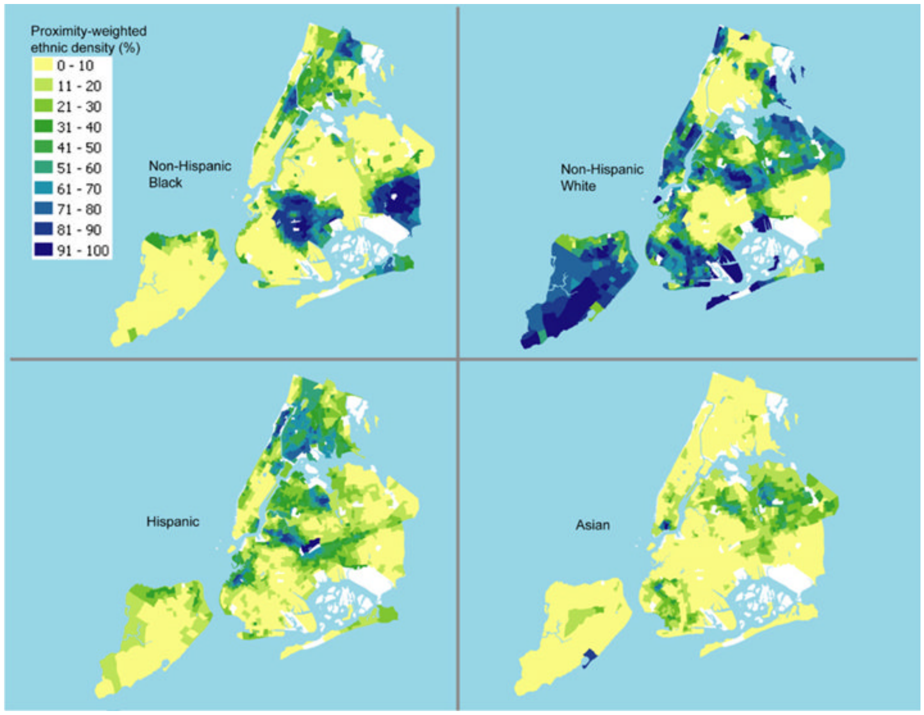

Fig. 3.
Geographic distribution, by census tract, of four ethnic densities measured with a 500-meter radius: New York City, 2000 US Census.

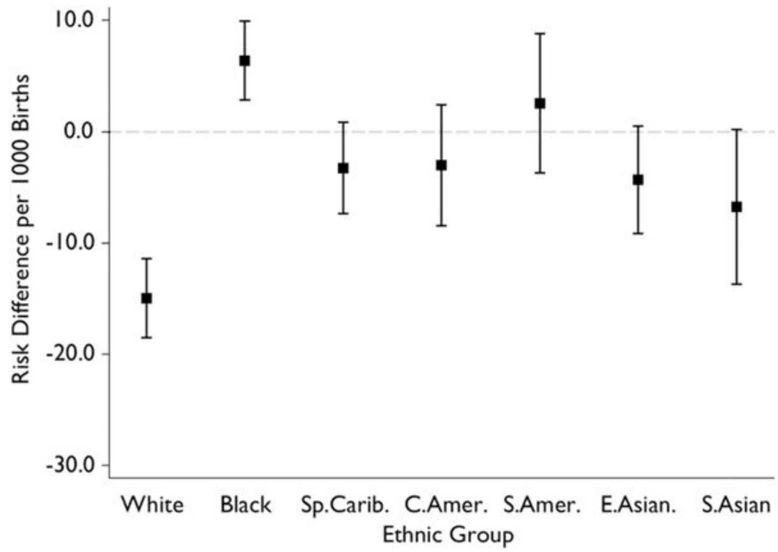

Fig. 4.

Adjusted risk differences with 95% confidence intervals for preterm birth among seven ethnic groups associated with residence in an ethnic enclave (> 25% ethnic density): New York City 1995–2003. Adjusted risk differences were calculated for US-born women aged 20–34 who were high-school educated, had 2–5 previous live births, were nonsmokers, received Medicaid, and resided in a more stable and poorer neighborhood. White=non-Hispanic white; Black=non-Hispanic black; Sp.Carib=Spanish Caribbean; C.Amer=Central American; S.Amer=South American; E.Asian=East Asian; S.Asian=South Asian.

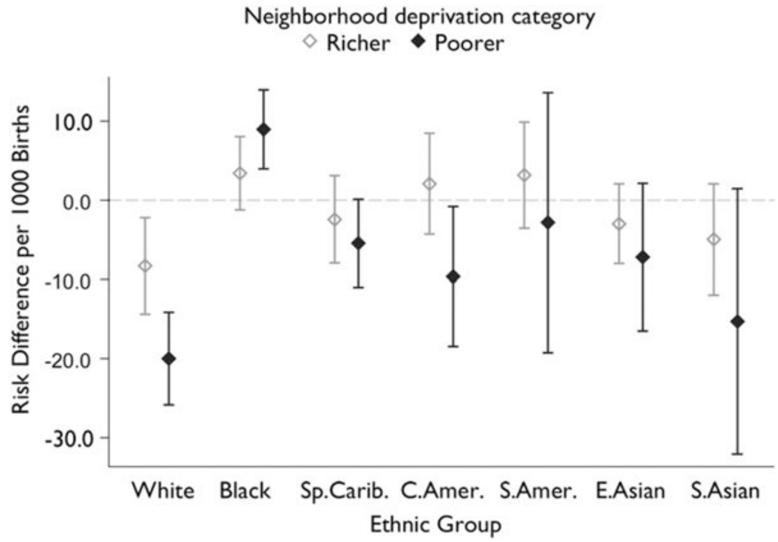

Fig. 5.

Neighborhood deprivation-stratified risk differences with 95% confidence intervals for preterm birth among seven ethnic groups associated with residence in an ethnic enclave (> 25% ethnic density): New York City 1995–2003. Risk differences were calculated for US-born women aged 20–34 who were high-school educated, had 2–5 previous live births, were nonsmokers, received Medicaid, and resided in a more stable neighborhood. White=non-Hispanic white; Black=non-Hispanic black; Sp.Carib=Spanish Caribbean; C.Amer=Central American; S.Amer=South American; E.Asian=East Asian; S.Asian=South Asian.

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Table 1a

Distribution of births to women in seven ethnic groups across covariate levels: New York City 1995–2003.

Variable	Ethnic Group													
	Non-Hispanic white		Non-Hispanic black		Spanish Caribbean		Central American		South American		East Asian		South Asian	
	N	%	*N*	%	*N*	%	*N*	%	*N*	%	*N*	%	*N*	%
Ethnic *N*, % of sample	237,528	27	256,673	29	181,539	20	73,096	8	43,324	5	53,867	6	41,860	5
Preterm Birth														
No	225,049	95	228,859	89	165,552	91	68,027	93	40,658	94	51,379	95	38,672	92
Yes	12,479	5	27,814	11	15,987	9	5069	7	2666	6	2488	5	3188	8
Ethnic density														
≤25%	23,813	10	40,657	16	30,959	17	18,939	26	11,428	26	24,573	46	29,063	69
>25%	213,715	90	216,016	84	150,580	83	54,157	74	31,896	74	29,294	54	12,797	31
Age (years)														
<20	5031	2	27,714	11	24,281	13	8349	11	2633	6	448	1	874	2
20–34	170,326	72	183,203	71	134,588	74	57,939	79	31,915	74	41,836	78	34,619	83
35+	62,171	26	45,756	18	22,670	12	6808	9	8776	20	11,583	22	6367	15
Maternal education (years)														
<12, age <20	2702	1	18,130	7	18,207	10	6532	9	1663	4	247	0	460	1
<12, age ≥20	15,526	7	46,091	18	49,101	27	36,667	51	10,619	25	13,274	25	8058	20
12	80,302	34	93,299	37	58,735	33	20,266	28	15,968	38	18,605	35	15,330	38
13–15	41,003	17	63,635	25	37,848	21	5382	8	8617	20	6395	12	6729	17
16+	95,868	41	31,678	13	15,504	9	2896	4	5657	13	14,345	27	9842	24
Previous births														
1	114,737	48	107,920	42	75,767	42	30,088	41	19,275	44	28,820	54	18,849	45
2–5	115,855	49	144,273	56	104,098	57	42,538	58	23,798	55	25,031	46	22,844	55
6+	6899	3	4469	2	1673	1	470	1	249	1	16	0	167	0
Prepregnancy weight (pounds)[a]														
<125	67,539	31	44,777	19	46,723	28	23,256	36	12,191	33	33,819	67	14,379	41
125–150	96,342	44	86,615	37	69,973	41	27,394	43	16,953	46	14,394	28	14,567	41
4150	55,268	25	104,617	44	51,984	31	13,376	21	7914	21	2432	5	6187	18

Variable	Ethnic Group													
	Non-Hispanic white		Non-Hispanic black		Spanish Caribbean		Central American		South American		East Asian		South Asian	
	N	%	*N*	%	*N*	%	*N*	%	*N*	%	*N*	%	*N*	%
Tobacco use														
Nonsmoker	228,429	96	240,397	94	171,968	95	72,572	100	42,920	99	53,477	100	41,672	100
Smoker	8436	4	14,690	6	8683	5	304	0	257	1	241	0	91	0
Late or no prenatal care[a]														
No	184,814	88	169,652	75	125,985	78	48,797	73	29,220	76	41,194	83	27,976	74
Yes	25,764	12	56,563	25	34,780	22	17,781	27	9455	24	8597	17	10,031	26
Payment for delivery														
Private insurance	170,109	72	86,774	35	47,773	26	7566	10	11,090	26	21,057	40	14,996	36
Medicaid	56,564	24	155,211	62	127,806	71	62,188	86	30,287	70	29,655	56	24,964	60
Self pay	9841	4	9095	4	4843	3	2826	4	1646	4	2539	5	1660	4
Nativity														
US-born	163,344	69	141,969	56	86,101	48	3685	5	4116	10	2490	5	950	2
Foreign-born	73,300	31	112,966	44	94,940	52	69,357	95	39,183	90	51,194	95	40,848	98
Residential stability														
Less stable	110,934	47	102,521	40	92,175	51	48,002	66	31,521	73	31,797	59	28,372	68
More stable	126,567	53	154,139	60	89,359	49	25,090	34	11,800	27	22,069	41	13,486	32
Neighborhood deprivation														
Richer	198,036	83	93,087	36	41,610	23	24,552	34	22,758	53	32,428	60	31,811	76
Poorer	39,458	17	163,559	64	139,911	77	48,538	66	20,563	47	21,437	40	10,044	24

[a]Variables were missing for less than 4% of observations, with exception of prepregnancy weight (missingness ranged from 6.0% in East Asians to 16.1% in South Asians) and prenatal care (missingness ranged from 7.6% in East Asians to 11.9% in non-Hispanic blacks).

Table 1b

Distribution of census tracts across neighborhood-level covariates: New York City 1995–2003.

Neighborhood-level covariate	***N***	**%**
Non-Hispanic white ethnic density		
≤25%	1084	50
>25%	1072	50
Non-Hispanic black ethnic density		
≤25%	1359	63
>25%	797	37
Hispanic ethnic density		
≤25%	1341	62
>25%	815	38
Asian ethnic density		
≤25%	1942	90
>25%	214	10
Residential stability		
Less stable	1015	47
More stable	1134	53
Neighborhood deprivation		
Richer	1394	65
Poorer	750	35

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Table 2

Difference in preterm birth risk (per 1000 births) associated with maternal residence in an ethnic enclave (> 25% ethnic group) for seven ethnic groups: New York City, 1995–2003.

Ethnic Group	Model			
	Crude	Adjusted	Stratified: Richer Neighborhoods	Stratified: Poorer Neighborhoods
	RD[a] (95% CI)	RD[a] (95% CI)	RD[a] (95% CI)	RD[a] (95% CI)
Non-Hispanic white	−17.0 (−20.9,−13.1)	−15.0 (−18.5,−11.4)	−8.3 (−14.4,−2.2)	−20.0 (−25.9,−14.1)
Non-Hispanic black	9.5 (6.0,13.1)	6.4 (2.8,9.9)	3.4 (−1.2,8.1)	9.0 (4.0,14.0)
Spanish Caribbean	−3.6 (−7.4,0.2)	−3.3 (−7.4,0.8)	−2.4 (−7.9,3.1)	−5.4 (−10.9,0.1)
Central American	−3.2 (−7.9,1.5)	−3.0 (−8.5,2.4)	2.1 (−4.2,8.5)	−9.6 (−18.5,−0.8)
South American	1.0 (−4.1,6.1)	2.5 (−3.7,8.8)	3.2 (−3.5,9.9)	−2.8 (−19.3,13.6)
East Asian	−3.7 (−7.4,−0.1)	−4.3 (−9.1,0.5)	−3.0 (−8.0,2.1)	−7.2 (−16.5,2.1)
South Asian	−9.3 (−16.0,−2.6)	−6.7 (−13.7,0.2)	−4.9 (−11.9,2.1)	−15.3 (−32.0,1.4)

[a] RD=risk difference; adjusted and stratified RDs were calculated for US-born women aged 20–34 who were high-school educated, had 2–5 previous live births, were nonsmokers, received Medicaid, and resided in a more stable and (for adjusted estimates) poorer neighborhood.