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Neighborhood Deprivation and Privilege: an Examination of Racialized-Economic Segregation and Preterm Birth, Florida 2019

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Abstract

The Black-White disparity in preterm birth persists and is not fully explained by individual-level social, behavioral, or clinical risk factors. Consequently, there is increasing emphasis on understanding the role of structural and area-level factors. Racialized-economic segregation measured as the index of concentration at the extremes (ICE) simultaneously captures extremes of deprivation and privilege. Our objective was to examine associations between preterm birth (PTB) and the index of concentration at the extremes (ICE). In this cross-sectional study, we analyzed 193,957 Florida birth records from 2019 linked to 2015–2019 census tract data from the American Community Survey. We assessed PTB (< 37 weeks gestation) by subtypes: (1) early (< 34 weeks) and late (34–36 weeks) and (2) spontaneous and indicated (i.e., provider-initiated) deliveries. We calculated adjusted odds ratios (aOR) and 95% confidence intervals (CI) for three ICE measures: (1) ICE_INC: income, (2) INC_INC + WB: income + race/ethnicity (non-Hispanic White vs. Black), and (3) INC_INC + WH: income + race/ethnicity (non-Hispanic White vs. Hispanic). Results. For ICE_INC and INC_INC + WB, aORs for residing in the worst-off vs. best-off areas were 1.25 (95% *CI*: 1.12, 1.46) and 1.21 (95% *CI*: 1.07, 1.37) for early PTB, respectively, and 1.16 (95% *CI*: 1.05, 1.28) to 1.22 (95% *CI*: 1.12, 1.34) for indicated PTB. In conclusion, deprivation captured by ICE was associated with increased odds of early or indicated

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Ethics Approval The Ethics and Human Research Protection Program at the Florida Department of Health deemed this activity as a public health practice, not research involving human subjects. No ethical approval is required.

Competing Interests The authors declare no competing interests.

Disclaimer The findings and conclusions in this report are those of the authors and do not necessarily represent the official position of the US Centers for Disease Control and Prevention or the Florida Department of Health.

PTB. Eliminating PTB disparities may require a multifaceted approach that includes addressing the interplay between income and race/ethnicity in residential areas.

Keywords

Preterm birth; Index of concentration at the extremes; ICE; Neighborhood poverty; Deprivation

Background

Preterm birth (< 37 completed weeks gestation) is a major risk factor for infant mortality and other adverse outcomes with serious health implications [1] In Florida, 10.5% of live births in 2020 were preterm [2], similar to the US rate of 10.1% [3]. While preterm birth rates were comparable between Hispanic (9.2%) and non-Hispanic White infants (9.3%) in Florida, the rate was 1.5 times higher for non-Hispanic Black infants than non-Hispanic White infants [2].

This disparity, which is also observed nationwide [3] is not fully explained by individual-level social, behavioral, and clinical risk factors of preterm birth [4]. Consequently, there is increasing emphasis on understanding the role of structural and area-level factors. Factors such as racial residential segregation and neighborhood conditions (e.g., socioeconomic disadvantage, crime, and pollution) where mothers reside before or during pregnancy are independently associated with increased risk of preterm birth, even after adjusting for individual-level factors [1, 5].

However, racial segregation and income disparities may demonstrate a stronger negative impact on adverse birth outcomes when they interact than when assessed alone. Racialized-economic segregation or the index of concentration at the extremes (ICE) is a measure of spatial polarization that simultaneously captures extremes of privilege and deprivation [6]. Five studies [6–10] examined the association between ICE and preterm birth in California, Massachusetts, and New York. They found that women residing in the most economically deprived areas with high proportions of Black residents were more likely to experience preterm birth than women residing in the most economically privileged areas with high proportions of White residents. However, researchers did not assess ICE with respect to the distribution of Hispanic populations or compare categories of preterm birth subtypes, namely (1) spontaneous and indicated (i.e., provider-initiated) preterm delivery or (2) live births born at 24–33 or 34–36 week gestation. These preterm birth subtypes have different etiologies.

To address these gaps, we examined the association between maternal residence in areas of concentrations of low vs. high economic and racial privilege (as measured by ICE, including a measure for Hispanic ethnicity) and preterm birth, overall and by subtypes.

Methods

Data Source

We linked 2019 Florida birth certificate records to the 2015–2019 American Community Survey (ACS) 5-year estimates by census tract of maternal residence [11]. This linkage created a dataset containing individual-level measures and area-level variables based on census tract population estimates [12]. The Ethics and Human Research Protection Program at the Florida Department of Health deemed this activity as a public health practice, not research involving human subjects.

Measures

Our outcome, preterm birth, was based on the clinician's obstetric estimation of gestational age at birth [13]. We defined the percentage of preterm birth as the proportion of total live singleton births born < 37 completed weeks multiplied by 100. We also categorized preterm birth as early (24–33 weeks) and late (34–36 weeks) and according to delivery type (spontaneous or indicated) [14].

ICE variables capture spatial social and economic polarization through scores quantifying extremes of socioeconomic deprivation and privilege [15]. We categorized individuals into tertiles using these scores, ranging from − 1 to 1 for the following three ICE exposures: [16]

- Income (ICE_INC). The number of persons with high annual household income (80th percentile for cut-point ≥ $100,000)[17] minus the number of persons with low annual household income (20th percentile for cut-point < $20,000), divided by the total population with known household incomes in the census tract (ACS website table B19001);
- Income and race/ethnicity for non-Hispanic White vs. Black (ICE_INC + WB). The number of non-Hispanic White persons with high annual household income (≥ $100,000) (ACS website tables B19001H, B19001B) minus the number of Black persons with low annual household income (< $20,000), divided by the total population with known income in the census tract; and
- Income and race/ethnicity for non-Hispanic White vs. Hispanic (ICE_INC + WH). The number of non-Hispanic White persons with high annual household income (≥ $100,000) (ACS website tables B19001H, B19001I) minus the number of Hispanic persons with low annual household income (< $20,000), divided by the total population with known income in the census tract.

We selected individual-level factors to add to statistical models a priori and excluded factors that were potential intermediates on the causal pathway (e.g., chronic diseases and prepregnancy body mass index). The analysis included maternal age (5-year increments from 15 through 44 years) and maternal race/ethnicity (Hispanic, non-Hispanic Black, and non-Hispanic White) [1]. For sensitivity analyses, we assessed paternal race/ethnicity (Hispanic, non-Hispanic Black, and non-Hispanic White) and maternal education (< high school (HS), HS diploma/GED, some college/associate degree, ≥ bachelor's degree).

For area-level factors added to statistical models, we examined (1) poverty rates to assess whether racialized economic segregation had an independent effect beyond that of poverty and (2) location (latitude and longitude coordinates) to account for maternal residences that were not spatially independent otherwise known as the distance-decay relationship [18]. For poverty rates, we divided the number of persons below 200% of the federal poverty level by the total population in the census tract (ACS website table B17001) and multiplied by 100. We grouped these percentages into tertiles where the cut-point for Tertile 1 was 19.7% (low poverty census tract) and for tertile 3 was 57.0% (high poverty census tract). Latitude and longitude, variables found in the linked birth file, were kept continuous and are based on the centroid of the residential census tract.

Study Population and Exclusions

Our study population included 201,602 infants born in Florida with resident mothers aged 15–44 years who were Hispanic, non-Hispanic Black, or non-Hispanic White. We excluded multiple gestation pregnancies (e.g., twins) ($n = 6226$) because these births have a higher risk of preterm birth than singletons due to spontaneous labor or premature rupture of membranes (PROM), or as a result of maternal conditions such as pre-eclampsia or fetal disorders [19]. In addition, we excluded gestational ages < 24 to maximize the inclusion of viable births [20] and > 41 weeks because these infants have different morbidity and mortality risks than term infants ($n = 885$) [21]; and records missing data on the delivery type ($n = 524$) or ICE variables ($n = 10$), leaving an analytic sample of 193,957. We excluded births after 41-week gestation (postterm) because they are at increased risk of stillbirth or neonatal death than infants born at term 37–41 weeks. [22] Comparing preterm births to postterm births may dilute our findings because racial residential segregation is also associated with an increased risk of stillbirth [23]. Compared with the sample before implementing the exclusion criteria, there were no differences observed for maternal age (mean 29 vs. 29 years) or maternal race/ethnicity (non-Hispanic White (44% vs. 44%), non-Hispanic Black (23% vs. 22%), and Hispanic (33% vs. 33%) women).

Statistical Analysis

We examined the distribution of preterm birth (overall and subtypes) by individual-level and area-level characteristics using chi-square statistics. Odds ratios (OR) and 95% confidence intervals (CI) for preterm birth overall in relation to three ICE variables were estimated using random effects logistic regression models to account for clustering at the census tract level. Term births (37–41 weeks) ($n = 177{,}056$) served as the referent category for the four binary preterm birth subtype variables: (1) two variables for the gestational age at birth sub-analyses ($n = 181{,}226$ for < 34 weeks gestation and $n = 189{,}787$ for 34–36 weeks gestation) and (2) two variables for the delivery type sub-analyses ($n = 185{,}760$ for spontaneous preterm birth and $n = 185{,}253$ for indicated preterm birth). For these four preterm subtype variables, we used random effects logistic regression models because our random effects multinomial logistic regression models did not converge. We plotted fully adjusted odds ratios and 95% *CIs* using a logarithmic scale. The intraclass correlation coefficient for all random effect models was assessed. Odds ratios and 95% *CIs* from final models were adjusted for maternal age, maternal race/ethnicity, and area-level variables (i.e., poverty rate, and latitude and longitude coordinates). Each model included one of three ICE variables to avoid collinearity. Model 1 included one ICE variable alone. Model 2 included an ICE variable plus all individual-level variables. Model 3 included an ICE variable and all individual and area-level variables. To test for multicollinearity, we calculated the variance inflation factor, which had a value of 2.01 for our full model. Because this value was less than five, we had little concern that our model returned unreliable coefficient estimates [24, 25]. We conducted a sensitivity analysis to additionally adjust for maternal education and paternal race/ethnicity in full models. Statistical significance was considered at alpha < 0.05. All analyses were conducted in STATA v.17.0.

Results

Our analytic sample included 193,957 birth records and 4116 census tracts. The mean value was − 0.002 (standard deviation(SD) = 0.24) for ICE_INC (i.e., in the direction of a more extreme concentration of census tracts with lower median household incomes); 0.29 (SD = 0.44) for ICE_INC + WB (i.e., in the direction of a more extreme concentration of non-Hispanic White residents with high-income than Black residents with low-income); and 0.20 (SD = 0.47) for ICE_INC + WH (i.e., in the direction of a more extreme concentration of non-Hispanic White residents with high-income than Hispanic residents with low-income).

Approximately, 10% of births were preterm (Table 1). Compared with births delivered 37–41-week gestation, higher proportions of women with preterm births were 35–44 years of age, non-Hispanic Black, resided in high-poverty census tracts, and were grouped in tertile 1 for all ICE measures. Analyses of preterm birth subtypes yielded similar results except for maternal age: older women had a higher prevalence of indicated preterm births (26.3%) than term births (18.6%) whereas teenage mothers had a higher prevalence of spontaneous preterm births (5.9%) than term births (4.5%). All comparisons were statistically significant at $p < 0.001$.

Unadjusted odds ratios for preterm birth overall and ICE variables for levels of extreme deprivation and segregation (tertile 1) ranged from 1.34 to 1.47 and were statistically significant (Table 2). These associations attenuated after adjustment for individual-level factors (Model 2). Estimates for all ICE measures were still statistically significant. Although additional adjustment for area-level variables (Model 3) further weakened associations with ICE, women who lived in areas with both low household incomes and high proportions of Black residents experienced 13% increased odds (aOR = 1.13, 95% *CI*: 1.06, 1.20) of preterm birth than women who lived in areas with high household incomes and high proportions of White residents. For Hispanic populations, however, there was no longer a statistically significant association between ICE and overall preterm birth in the fully adjusted model (aOR = 1.05, 95% *CI*: 0.99, 1.12). Across the fully adjusted models, the intraclass correlation coefficient was very modest (0.004).

In sensitivity analyses, we observed minor changes in effect estimates after adding maternal education and paternal race/ethnicity to the models. Adjusted odds ratios were 1.06 (1.00, 1.13) for tertile 1 of ICE_INC, 1.10 (1.03, 1.18) for tertile 1 of ICE_INC + WB, and 1.03 (0.96, 1.10) for tertile 1 of ICE_INC + WH (table not shown).

Figure 1 shows fully adjusted odds ratios, plotted on a logarithmic scale, with tertiles of the three ICE measures for preterm birth subtypes: gestational timing and delivery type. No subtype variable was associated with the ICE measure comparing census tracts with mostly Hispanic residents and low household incomes to census tracts with mostly non-Hispanic White residents and high household incomes. The remaining two ICE variables (tertile 1) were consistently associated with increased odds of early preterm (24–33 weeks) and indicated preterm birth but not late preterm (34–36 weeks) or spontaneous preterm birth. For example, tertile 1 (worst-off) of ICE_INC was associated with 25% higher odds of early preterm birth than tertile 3 (best-off) of ICE_INC (aOR = 1.25, 95% *CI*: 1.12, 1.46). Also, relative to mothers living in areas with residents who are mostly non-Hispanic White and have high household incomes, mothers living in areas with residents who are mostly Black and have low household incomes (ICE_INC + WB) were 22% more likely to have an indicated preterm birth (aOR = 1.22, 95% *CI*: 1.12, 1.34). The supplemental figure shows the association by subtype when controlling for individual factors only. Associations were significant and patterns changed when controlling for area-level factors (ESM Fig. 1).

Discussion

In our study, most ICE measures were associated with preterm birth overall as well as the indicated and early preterm birth subtypes, after adjustment of individual and area-level measures. The ICE measure that did not remain associated with preterm birth after adjustment compared census tracts with low household income and high Hispanic populations to census tracts with high household income and high non-Hispanic White populations. Although adjusted associations between overall preterm birth and the lowest ICE tertile (representing the worst-off census tracts) weakened after controlling for area-level factors, particularly poverty rate, associations for two of three ICE variables (ICE_INC and ICE_INC + WB) remained statistically significant.

The poverty rate is an absolute measure that is based on how many persons in a defined geographic area are below the federal poverty threshold. ICE_INC, for example, is a relative measure that is based on the distribution between poor and wealthy individuals. The statistically significant association between ICE_INC + WB and preterm birth that remained after adjusting for poverty may indicate that there is a component of racialized-economic segregation (ICE) that had an impact on preterm birth, beyond the effect of poverty (e.g., lack of access to supermarkets, health services, sound housing stock, and highly ranked schools) [8]. In our population, after adjusting for poverty, there was no longer an adverse association between ICE_INC + WH and preterm birth. This suggests that, perhaps, the spurious association between ICE_INC + WH and preterm birth observed before adjustment acted through circumstances of poverty. Once we removed the effect of our poverty measure, preterm birth rates were similar between mothers residing in areas with mostly Hispanic residents and low household incomes to mothers residing in areas with mostly non-Hispanic White residents and high household incomes. This is consistent with the phenomenon where, relative to non-Hispanic US White mothers, Hispanic mothers are more likely to live in socioeconomically deprived areas but have similar rates of adverse birth outcomes [26]. The paradox is often attributed to social cohesion and cultural norms,[26, 27] which may have buffered away any remaining effect of racialized-economic segregation after controlling for poverty in our study. Nonetheless, these are speculations, and future research to understand the differences between ICE_INC + WH and the other ICE measures is warranted.

The effect of poverty may act through different mechanisms. Parker and colleagues [28] found that, compared with 56% of families with term infants, 61% of families with preterm infants experienced at least one unmet need of health care hardship (e.g., forgoing needed medical care, dental care, or prescriptions due to inability to pay), energy insecurity (e.g., a threatened or actual shut-off of gas or electricity because of non-payment), housing instability (e.g., moving two or more times in the past year), or household food insecurity (e.g., could not afford to eat balanced meals). The authors suggest that early screening for unmet needs, particularly during the prenatal period, followed by referral to appropriate services could reduce the risk of preterm birth.

Additionally, other outcomes related to preterm birth have been shown to benefit from social service interventions. Specifically, infant mortality was positively affected by income support (earned income tax credit), and low birthweight by care coordination and community outreach (participation in the Healthy Start Program) and nutrition support (participation in the Special Supplemental Nutrition Program for Women, Infants, and Children) [29].

Our findings for overall preterm birth are comparable with those from previous studies. Two studies presented only crude estimates for ICE exposures experienced during adulthood [6, 10] and two controlled for individual-level factors alone [7, 8]. Relative risks reported in these four studies ranged from 1.14 to 1.29 for ICE: income and 1.17 to 1.36 for ICE: income + race. Krieger et al.[9] controlled for Home Owner's Loan Corporation (HOLC) grade (redlining), a measure of structural racism, in addition to maternal factors. They reported residing in New York City census tracts with predominantly low-income Black households was associated with 25% increased odds of preterm birth compared with residing in areas with high-income White households (aOR = 1.25; 95% *CI*: 1.20, 1.30).

Huynh et al. [8] examined poverty as a separate variable and noted high neighborhood poverty was associated with increased odds of preterm birth (aOR = 1.09; 95% *CI* 1.04, 1.14). They did not examine whether the association with poverty persisted after accounting for racialized-economic segregation as measured by ICE indictors. We found no studies on preterm birth that assessed ICE indicators according to the distribution of Hispanic populations. Comparable to our early preterm (< 34-week gestation) ICE analyses, Shrimali et al. [10] found the risk of birth < 32-week gestation or "very preterm birth" had larger estimates and less precision (unadjusted OR quintile 1 vs. 5 = 1.60; 95% *CI* 1.46, 1.75) than the "overall preterm" < 37-completed-week gestation outcome (unadjusted OR quintile 1 vs. 5 = 1.30; 95% *CI* 1.26, 1.34). They did not present data on late preterm births.

The etiology of preterm birth is multifactorial [30] and can be differentiated to some extent by examining groups based on gestational age at birth or reasons for preterm deliveries that are indicated. Infants born before 34 weeks completed gestation have higher morbidity and mortality than infants born at later gestational ages, which is likely related to inadequate fetal lung maturation often requiring intensive care [31]. In addition, preterm births delivered spontaneously following a premature rupture of membranes or preterm labor have a different etiology than preterm birth deliveries that are indicated due to maternal and fetal complications necessitating delivery (e.g. hyper-tensive disorders of pregnancy, gestational diabetes, and fetal distress) [14, 32]. Moreover, indicated preterm births have higher mortality than spontaneous preterm births [33]. Janevic et al. recently demonstrated that more than one-third of the association between racialized-economic segregation and the morbidity and mortality of very preterm neonates is explained by the location of delivery hospitals [34]. That women in low-income majority Black areas are more likely to give birth at local resource-limited hospitals with lower quality of care than at hospitals in more privileged areas. Therefore, understanding the different etiologies of preterm birth can support the prioritization of prevention strategies for those at highest risk [1, 35, 36], which in turn may help to mitigate the observed effect of racialized-economic segregation on preterm birth. The use of our findings to identify new interventions and policies that can weaken this association is an opportunity for future research.

A limitation of our study is the use of census tract variables to represent maternal residential conditions during pregnancy and shortly after birth. Despite the dual dimensionality of ICE as a measure, neighborhood conditions are multi-dimensional making it difficult to capture complex neighborhood conditions in one measure. Also, we did not adjust for residential mobility before or during pregnancy. Studies found that while 12–33% of participants moved during pregnancy [37, 38], many typically moved within the same municipality [38]. We are also limited to data collected through birth records and therefore are unable to account for factors such as individual maternal income, potentially yielding residual confounding. Regarding the algorithm used to classify spontaneously and indicated preterm birth, Klebanoff et al. [14] reported inter-rater disagreement between 5 and 15% and explained that the errors of commission and omission of diagnoses in the birth certificate that they found would yield erroneous results from the algorithm. This misclassification of outcome would be non-differential and bias our estimates toward the null. Lastly, we did not examine the association in other populations with higher rates of infant mortality and racial segregation (e.g., American Indian/Alaska Natives). This is an area of future research.

Our study has many strengths. Past analyses have been based on data from New York, California, and Massachusetts, but none from Florida, a Southeastern state. Thus, analyzing data from the Southern region of the USA, where preterm birth rates are relatively high [39], may help to identify nuances that allow public health and clinical interventions to strategically focus resources on groups most affected by racial and economic disparities. Also, we analyzed data from a large and diverse population that allowed for the calculation of robust estimates and an ICE measure specific to the ethnic and economic experience of Black and Hispanic populations in Florida. Our measures, however, did not capture specific populations within Black and Hispanic communities in Florida (e.g., Cubans, Haitians, and other residents with familial ties to the nearby Caribbean region) who may have diverse neighborhood social environments. This may be an area for future research as the 2020 Census will soon produce data on more detailed racial and ethnic groups [40].

Public Health Implications

Population-based studies that apply multilevel modeling techniques to clinical data could help elucidate the effect of social determinants of health (e.g., measures of socioeconomic deprivation and racial privilege) on preterm birth [36]. This understanding could help inform public health strategies and social policies to improve maternal and infant health, particularly in socially deprived communities at high risk for preterm birth [1, 35, 36].

Conclusions

In this study, compared with the best-off census tracts, the worst-off census tracts for our ICE_INC and INC_INC + WB measures were associated with increased odds of overall, early, and indicated preterm, after adjustment of individual and area-level factors, including poverty rate. While our study examined neighborhood measures jointly to capture multidimensional impact, specifically that of race/ethnicity and income, more research could be done to examine the mechanisms behind the effect of ICE on preterm birth and its subtypes.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

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Data Availability

The data that support the findings of this study are from the Florida Department of Health and are not publicly available due to the sensitive identifiable data, which we used under license for the current study.

References

1. Institute of Medicine Committee on Understanding Premature B, Assuring Healthy O. The national academies collection: reports funded by National Institutes of Health. In: Behrman RE, Butler AS, editors. Preterm Birth: Causes, Consequences, and Prevention. National Academies Press (US) Copyright © 2007, National Academy of Sciences; 2007.
2. Florida Birth Certificate Data. Accessed October 31, 2021. https://www.flhealthcharts.gov/
3. Martin JA, Hamilton BE, Osterman MJK. Births in the United States, 2020. Vol. 418. 2021. NCHS Data Brief.
4. Thoma ME, Drew LB, Hirai AH, Kim TY, Fenelon A, Shenassa ED. Black-white disparities in preterm birth: geographic, social, and health determinants. Am J Prev Med. 2019;57(5):675–86. 10.1016/j.amepre.2019.07.007. [PubMed: 31561920]
5. Mehra R, Boyd LM, Ickovics JR. Racial residential segregation and adverse birth outcomes: a systematic review and meta-analysis. Social science & medicine. 1982;191:237–50. 10.1016/j.socscimed.2017.09.018.
6. Krieger N, Waterman PD, Batra N, Murphy JS, Dooley DP, Shah SN. Measures of local segregation for monitoring health inequities by local health departments. Am J Public Health. 2017;107(6):903–6. 10.2105/ajph.2017.303713. [PubMed: 28426303]

7. Chambers BD, Baer RJ, McLemore MR, Jelliffe-Pawlowski LL. Using index of concentration at the extremes as indicators of structural racism to evaluate the association with preterm birth and infant mortality-California, 2011–2012. J Urban Health : Bulletin of New York Academy Med. 2019;96(2):159–70. 10.1007/s11524-018-0272-4.

8. Huynh M, Spasojevic J, Li W, et al. Spatial social polarization and birth outcomes: preterm birth and infant mortality - New York City, 2010–14. Scandinavian J Public Health. 2018;46(1):157–66. 10.1177/1403494817701566.

9. Krieger N, Van Wye G, Huynh M, et al. Structural racism, historical redlining, and risk of preterm birth in New York City, 2013–2017. Am J Public Health. 2020;110(7):1046–53. 10.2105/AJPH.2020.305656. [PubMed: 32437270]

10. Shrimali BP, Pearl M, Karasek D, Reid C, Abrams B, Mujahid M. Neighborhood privilege, preterm delivery, and related racial/ethnic disparities: an intergenerational application of the index of concentration at the extremes. Am J Epidemiol. 2020;189(5):412–21. 10.1093/aje/kwz279. [PubMed: 31909419]

11. Bureau USC. When to use 1-year or 5-year estimates. Accessed April 24, 2022, https://www.census.gov/programs-surveys/acs/guidance/estimtes.html

12. Rajaratnam JK, Burke JG, O'Campo P. Maternal and child health and neighborhood context: the selection and construction of area-level variables. Health Place. 2006;12(4):547–56. 10.1016/j.healthplace.2005.08.008. [PubMed: 16188483]

13. Martin JA, Osterman MJ, Kirmeyer SE, Gregory EC. Measuring gestational age in vital statistics data: transitioning to the obstetric estimate National vital statistics reports : from the Centers for Disease Control and Prevention, National Center for Health Statistics. National Vital Statistics System. 2015;64(5):1–20.

14. Klebanoff MA, Yossef-Salameh L, Latimer C, et al. Development and validation of an algorithm to determine spontaneous versus provider-initiated preterm birth in US vital records. Paediatr Perinat Epidemiol. 2016;30(2):134–40. 10.1111/ppe.12267. [PubMed: 26860444]

15. Krieger N, Waterman PD, Spasojevic J, Li W, Maduro G, Van Wye G. Public health monitoring of privilege and deprivation with the index of concentration at the extremes. Am J Public Health. 2016;106(2):256–63. 10.2105/ajph.2015.302955. [PubMed: 26691119]

16. Krieger N, Kim R, Feldman J, Waterman PD. Using the Index of concentration at the extremes at multiple geographical levels to monitor health inequities in an era of growing spatial social polarization: Massachusetts, USA (2010–14). Int J Epidemiol. 2018;47(3):788–819. 10.1093/ije/dyy004. [PubMed: 29522187]

17. Bureau USC. Historical Income Tables: Table H-1. Income limits for each fifth and top fifth (All Races). Accessed December 28, 2021, https://www.census.gov/data/tables/time-series/demo/income-poverty/historical-income-households.html

18. Morlon H, Chuyong G, Condit R, et al. A general framework for the distance-decay of similarity in ecological communities. Ecol Lett 2008;11(9):904–17. 10.1111/j.1461-0248.2008.01202.x. [PubMed: 18494792]

19. Blondel B, Kogan MD, Alexander GR, et al. The impact of the increasing number of multiple births on the rates of preterm birth and low birthweight: an international study. Am J Public Health. 2002;92(8):1323–30. 10.2105/ajph.92.8.1323. [PubMed: 12144992]

20. Stoll BJ, Hansen NI, Bell EF, et al. Neonatal outcomes of extremely preterm infants from the NICHD Neonatal Research Network. Pediatrics. 2010;126(3):443–56. 10.1542/peds.2009-2959. [PubMed: 20732945]

21. Practice bulletin no. 146: Management of late-term and postterm pregnancies. Obstet Gynecol. 2014;124(2 Pt 1):390–6. 10.1097/01.AOG.0000452744.06088.48. [PubMed: 25050770]

22. Middleton P, Shepherd E, Crowther CA. Induction of labour for improving birth outcomes for women at or beyond term. Cochrane Database Syst Rev. 2018;5(5):D004945. 10.1002/14651858.CD004945.pub4.

23. Williams AD, Wallace M, Nobles C, Mendola P. Racial residential segregation and racial disparities in stillbirth in the United States. Health Place. 2018;51:208–16. 10.1016/j.healthplace.2018.04.005. [PubMed: 29715639]

24. Kim JH. Multicollinearity and misleading statistical results. Korean J Anesthesiol. 2019;72(6):558–69. 10.4097/kja.19087. [PubMed: 31304696]
25. Belsley DA, Kuh E, Welsch RE. Regression diagnostics: identifying influential data and sources of collinearity. Wiley & Sons; 2005.
26. Franzini L, Ribble JC, Keddie AM. Understanding the Hispanic paradox. Ethn Dis Autumn. 2001;11(3):496–518.
27. Shaw RJ, Pickett KE. The health benefits of Hispanic communities for non-Hispanic mothers and infants: another Hispanic paradox. Am J Public Health. 2013;103(6):1052–7. 10.2105/AJPH.2012.300985. [PubMed: 23597369]
28. Parker MG, de Cuba SE, Rateau LJ, et al. Household unmet basic needs in the first 1000 days and preterm birth status. J Perinatol. 2022. 10.1038/s41372-022-01325-5.
29. Taylor LA, Tan AX, Coyle CE, et al. Leveraging the social determinants of health: what works? PLoS ONE. 2016;11(8):e0160217. 10.1371/journal.pone.0160217. [PubMed: 27532336]
30. Goldenberg RL, Culhane JF, Iams JD, Romero R. Epidemiology and causes of preterm birth. Lancet (London, England). 2008;371(9606):75–84. 10.1016/s0140-6736(08)60074-4. [PubMed: 18177778]
31. Cunningham FG, Leveno KJ, Bloom SL, Hauth JC, Rouse DJ, Spong C. Williams obstetrics. 23 ed. McGraw-Hill; 2010.
32. Ananth CV, Vintzileos AM. Maternal-fetal conditions necessitating a medical intervention resulting in preterm birth. Am J Obstet Gynecol. 2006;195(6):1557–63. 10.1016/j.ajog.2006.05.021. [PubMed: 17014813]
33. Kamath-Rayne BD, DeFranco EA, Chung E, Chen A. Subtypes of preterm birth and the risk of postneonatal death. J Pediatr. 2013;162(1):28–34.e2. 10.1016/j.jpeds.2012.06.051. [PubMed: 22878113]
34. Janevic T, Zeitlin J, Egorova NN, et al. Racial and economic neighborhood segregation, site of delivery, and morbidity and mortality in neonates born very preterm. J Pediatr. 2021;235:116–23. 10.1016/j.jpeds.2021.03.049. [PubMed: 33794221]
35. Committee on Practice Bulletins—Obstetrics The American College of Obstetricians and Gynecologists. Practice bulletin no 130: prediction and prevention of preterm birth. Obstetr Gynecol 2012;120(4):964–73. 10.1097/AOG.0b013e3182723b1b
36. Shapiro-Mendoza CK, Barfield WD, Henderson Z, et al. CDC grand rounds: public health strategies to prevent preterm birth. MMWR Morbidity and mortality weekly report. 2016;32:826–30. 10.15585/mmwr.mm6532a4.
37. Canfield MA, Ramadhani TA, Langlois PH, Waller DK. Residential mobility patterns and exposure misclassification in epidemiologic studies of birth defects. J Eposure Sci Environ Epidemiol. 2006;16(6):538–43. 10.1038/sj.jes.7500501.
38. Fell DB, Dodds L, King WD. Residential mobility during pregnancy. Paediatr Perinat Epidemiol. 2004;18(6):408–14. 10.1111/j.1365-3016.2004.00580.x. [PubMed: 15535816]
39. National Center for Health Statistics. Percentage of births born preterm by state. Updated December 28, 2021. https://www.cdc.gov/nchs/pressroom/sosmap/preterm_births/preterm.htm.
40. Castro-Rodriguez JA, Forno E, Rodriguez-Martinez CE, Celedón JC. Risk and protective factors for childhood asthma: what is the evidence? The Journal Allergy Clinical Immunology In Practice. 2016;4(6):1111–22. 10.1016/j.jaip.2016.05.003.

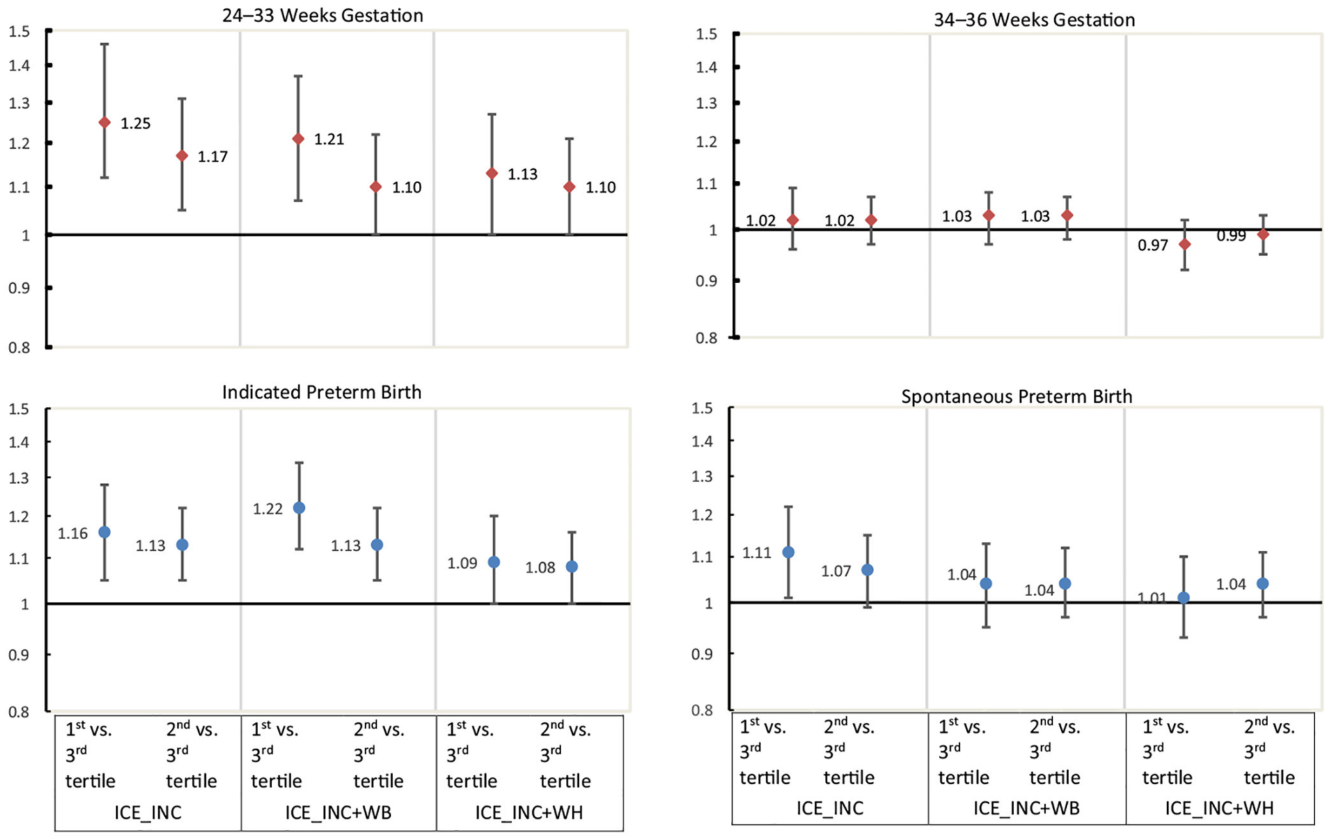

Fig. 1.

Adjusted odds ratios and 95% confidence intervals for preterm birth subtypes and three index of concentration at the extremes measures, Florida 2019. Abbreviations: ICE, index of concentration at the extremes; ICE_INC, low vs. high US household income; ICE_INC + WB = income and race/ethnicity for non-Hispanic White vs. Black; ICE_INC + WH, income and race/ethnicity for non-Hispanic White vs. Hispanic. All odds ratios adjusted for maternal age, maternal race/ethnicity, poverty, and latitude and longitude coordinates are plotted on a logarithmic scale

Abbreviations: ICE = index of concentration at the extremes; ICE_INC = low vs. high U.S. household income; ICE_INC+WB = income and race/ethnicity for non-Hispanic White vs. Black; ICE_INC+WH= income and race/ethnicity for non-Hispanic White vs. Hispanic. All odds ratios adjusted for maternal age, maternal race/ethnicity, poverty, and latitude and longitude coordinates are plotted on a logarithmic scale.

Table 1

Prevalence of study population characteristics and of preterm birth, overall and by subtypes, by characteristics, Florida 2019

			Preterm birth total	Preterm birth by gestational weeks		Preterm birth by delivery type	
Characteristics	**Total %**	**37–41 weeks**	**24–36 weeks**	**34–36 weeks**	**24–33 weeks**	**Indicated**	**Spontaneous**
Total (*N*)	193,957	177,056	16,901	12,731	4170	8197	8704
Maternal age (years)							
15–19	4.5	4.5	4.8	4.6	5.3	3.5	5.9
20–24	18.9	18.9	18.5	18.9	17.4	15.2	21.6
25–29	29.0	29.2	26.5	26.7	26.2	25.5	27.5
30–34	28.7	28.8	27.9	27.7	28.5	29.5	26.4
35–44	19.0	18.6	22.3	22.2	22.6	26.3	18.5
Maternal race/ethnicity							
Hispanic	33.4	33.6	31.2	31.5	30.2	30.7	31.6
Non-Hispanic Black	22.4	21.6	31.4	29.3	37.8	32.6	30.2
Non-Hispanic White	44.2	44.9	37.5	39.2	32.0	36.4	38.2
Poverty (census tract)							
Tertile (T)3 (high poverty)	33.7	33.1	39.8	38.8	42.8	38.9	40.6
T2	33.4	33.5	32.7	32.9	32.3	32.6	32.8
T1 (low poverty)	32.9	33.4	27.5	28.4	24.9	28.5	26.6
ICE: income (low vs high)*							
T1 (least income)	33.8	33.3	39.6	38.5	42.7	38.7	40.4
T2	33.4	33.5	33.2	33.3	32.8	33.2	33.2
T3 (most income)	32.8	33.3	27.3	28.2	24.5	28.1	26.5
ICE: race/ethnicity + income (low-income Black vs high-income White non-Hispanic)†							
T1 (most Black and low-income)	33.3	32.7	39.5	38.1	44.0	40.0	39.1
T2	33.8	33.9	32.9	33.4	31.5	32.4	33.5
T3 (most White and high-income)	32.9	33.4	27.5	28.5	24.6	27.6	27.4
ICE: race/ethnicity + income (low-income Hispanic vs high-income White non-Hispanic)‡							
T1 (most Hispanic and low-income)	33.3	32.9	37.1	35.9	40.7	37.6	36.6
T2	33.7	33.6	34.7	34.8	34.4	34.1	35.4
T3 (most White and high-income)	33.0	33.5	28.2	29.3	24.9	28.4	28.0

Abbreviations: *ICE*, index of concentration at the extremes; *T*, tertile

All chi-square *p*-values comparing distributions for ≥ 37 weeks gestation with preterm births (overall and subtypes) by individual and area-level characteristics equaled < 0.001. Except for the first row, column percentages are presented and may not total 100% due to rounding

[*] Calculation: (# persons with high annual household income (≥ $100,000)) − (# persons with low annual household income (< $20,000)) / total population

[†] Calculation: (# non-Hispanic White persons with high annual household income (≥ $100,000)) − (# Black persons with low annual household income (< $20,000)) / total population

[‡] Calculation: (# non-Hispanic White persons with high annual household income (≥ $100,000)) − (# Hispanic persons with low annual household income (< $20,000)) / total population

Table 2

Odds ratios and 95% confidence intervals for preterm birth and measures of index of concentration at the extremes, Florida 2019

		Model 1*		Model 2†		Model 3‡	
Index of concentration at the extremes (ICE)	**% Preterm birth**	***OR***	**95% *CI***	***OR***	**95% *CI***	**OR**	**95% *CI***
ICE: income (low vs high)							
Tertile (T)1 (least income)	10.2	1.45	1.39, 1.51	1.33	1.27, 1.38	1.13	1.06, 1.21
T2	8.6	1.21	1.16, 1.27	1.18	1.13, 1.23	1.10	1.04, 1.16
T3 (most income)	7.3	1.00	Ref	1.00	Ref	1.00	Ref
ICE: race/ethnicity + income (low-income Black vs high-income White non-Hispanic)							
T1 (most Black and low-income)	10.4	1.47	1.41, 1.54	1.28	1.22, 1.34	1.13	1.06, 1.20
T2	8.5	1.18	1.13, 1.24	1.16	1.11, 1.22	1.09	1.03, 1.14
T3 (most White and high-income)	7.3	1.00	Ref	1.00	Ref	1.00	Ref
ICE: race/ethnicity + income (low-income Hispanic vs high-income White non-Hispanic)							
T1 (most Hispanic and low-income)	9.7	1.34	1.28, 1.40	1.18	1.13, 1.24	1.05	0.99, 1.12
T2	9.0	1.23	1.18, 1.28	1.16	1.11, 1.22	1.06	1.00, 1.11
T3 (most White and high-income)	7.4	1.00	Ref	1.00	Ref	1.00	Ref

Abbreviations: *CI*, confidence interval; *ICE*, index of concentration at the extremes; *OR*, odds ratio; *T*, tertile

* Model 1: Crude

† Model 2: Model 1 + individual-level characteristics (i.e., maternal age and race/ethnicity)

‡ Model 3: Model 2 + area-level characteristics (i.e., poverty, latitude, and longitude coordinates)