Structural Heteropatriarchy and Birth Outcomes in the United States

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ABSTRACT Emerging evidence links structural sexism and structural discrimination against lesbian, gay, and bisexual (LGB) populations to poor health outcomes, but studies have yet to examine the combined effects of these mutually reinforcing systems of inequality. Therefore, we developed a composite measure of structural heteropatriarchy—which includes state-level LGB policies, family planning policies, and indicators of structural sexism (e.g., women's political and economic position relative to men)—and examined its relationship to birth outcomes using data from Waves I to V of the National Longitudinal Study of Adolescent to Adult Health. Multivariate regression analyses demonstrated that higher levels of heteropatriarchy were associated with an increased risk of preterm birth and decreased birth weight, net of important covariates. There was no association between clinical low birth weight and heteropatriarchy, or interactions between heteropatriarchy and individuals' race, ethnicity or sexual identity, suggesting a negative effect of heteropatriarchy on birth outcomes for all pregnant people. This study demonstrates the importance of considering gender and sexuality as mutually reinforcing systems of oppression that impact population health. Future research should examine the impact of heteropatriarchy on additional health outcomes and in conjunction with other structural inequalities such as racism and transgender oppression.

KEYWORDS Gender • Sexuality • Structural stigma • Heteropatriarchy • Birth outcomes

Introduction

The relationship between discrimination—including that based on race, sexual orientation, and gender—and individual health outcomes is well documented (Ayhan et al. 2020; Krieger 2014; Williams and Mohammed 2013). Much of this research, however, has focused on interpersonal experiences with discrimination (for a review, see Krieger 2014). Researchers have built on social ecological theories of health (McLeroy et al. 1988) and expanded this area of study by documenting how discrimination at a *structural* level can impact health outcomes. These include recent studies that have advanced the theory of structural sexism (Homan 2019) and structural lesbian, gay, and bisexual (LGB) stigma

(Hatzenbuehler 2014). In this article, we expand on these concepts and argue that forms of structural discrimination that target women, as well as sexual and gender minorities, are rooted in a system of oppression designed to reproduce and reinforce the dominance of heterosexual cisgender men. This synergistic structural discrimination—which we call heteropatriarchy—may lead to adverse health outcomes.

This study explores the role of heteropatriarchy and its implications for birth outcomes, specifically birth weight and preterm birth. Using the National Longitudinal Study of Adolescent to Adult Health, we created a measure of heteropatriarchy comprising multiple state- and county-level indicators that capture economic, cultural, and political dimensions of heteropatriarchy. We focused on birth outcomes for two primary reasons. First, the United States lags far behind other developed countries in terms of its birth outcomes (March of Dimes 2017; March of Dimes et al. 2012), and studies have shown that important regional variations in birth outcomes exist within the nation (Brown et al. 2019; Brown Speights et al. 2017). Hence, understanding how social climates, including modifiable policy indictors, impact birth outcomes is of notable public health importance. Second, our measure of heteropatriarchy is likely to reveal a disproportionate impact on people with the capacity to get pregnant (e.g., women, transgender men, nonbinary persons), and previous research has established that birth outcomes are sensitive to discriminatory experiences, both interpersonal and structural (Krieger et al. 2013; Novak et al. 2017; Samari et al. 2020).

We found that as structural heteropatriarchy increases, so does the risk of adverse birth outcomes for infants. These findings contribute to a growing body of work showing how social contexts affect individual health, as well as the ways pregnant people's bodies function as sites of inequality with long-term implications for intergenerational health.

Background

Gender, Sexuality, and Systemic Oppression

Feminist theorists have long argued that gender and sexuality are inherently linked systems of oppression that seek to enforce a gendered social structure that situated women within the home, advocated for compulsory motherhood, and privileged men's access to power and resources (Warner 1991). Traditional gendered divisions of labor are therefore rooted in the enforcement of heterosexual, monogamous relationships: the importance of *both gender and sexuality* in maintaining systematic gender oppression is well established in feminist theory (Butler 1991; Rich 1980; Rubin 1975, 1998). Policies that punish homosexuality and enforce heterosexuality are inherently linked to, if not grounded in, ideologies that distinguish between men's and women's social and economic roles. Given the inherent links between gender and sexuality, these two systems reinforce one another and shape men's and women's roles in both the workforce and the home; these cannot operate independently of each other (Jackson 2006; Rich 1980). Consequently, the subjugation of women requires the marginalization of other nonheterosexual identities to maintain its legitimacy. Some theorists have argued that sexuality is the "linchpin" of gender inequality (MacKinnon 1982:533).

Normative heterosexuality and normative gender are inherently tied; these two systems reinforce each other through social norms and co-constitutive structures and

policies that create and sustain patriarchal social arrangements and influence the types of social, cultural, and economic capital available to individuals. Similar to gender, heterosexuality is an institution that operates at multiple levels. And while there have been shifts in how heterosexuality is codified nationally (e.g., the changing legality of same-sex marriage), multiple state-level policies persist that provide or deny access to resources and protections on the basis of sexuality. While these policies may not always directly impact health outcomes, they reinforce norms related to heterosexuality and gender, which both codify and signal to citizens their value in civil society.

Despite the longstanding theoretical linkage of gender and sexuality in terms of policies, context, and norms, these concepts have remained largely independent in population health research. The idea that researchers should take a broader approach to understanding patriarchal systems in research is not new; Ingraham (1994) argued that the lack of explicit examination of the role of heterosexuality in gender research not only renders heterosexuality as an invisible feature of patriarchal society, but also a *natural* one, thereby closing off critical engagement around how these two systems are bound together and are mutually constitutive.

While multiple terms have been used to articulate how sexuality functions in patriarchal society (e.g., heterogender (Ingraham 1994); heteronormativity (Warner 1991); heterosexism), in this article we employ the term "structural heteropatriarchy" to describe the system of regulation that privileges cisgender men and heterosexuality. We do so for several reasons. The term heteropatriarchy has been increasingly used in qualitative and theoretical research to highlight the interconnected nature of oppression based on gender (i.e., sexism) and on sexuality (i.e., heterosexism) (Arvin et al. 2013; Darwin 2018; Strolovitch et al. 2017; Valdes 1996).

Moreover, patriarchy, as a term, highlights how male dominance is central to the organization of a capitalist society, including laws, policies, and our educational and professional systems. Heterosexuality is an essential feature of patriarchal society. We therefore follow and expand on Ingraham's work (1994) to argue that the term "heteropatriarchy" allows us to explicitly name and interrogate the role of heterosexuality in the production of inequities. Thus, while some terms acknowledge the interconnectedness of sexuality and gender, the term heteropatriarchy focuses on not just how individuals with marginalized identities may be targeted by individual persons, but also on the interlocking *systems* of oppression that operate on multiple levels to privilege both men and heterosexuality.

We argue that structural sexism and structural LGB stigma are inherently linked systems of oppression that co-constitute structural heteropatriarchy because they serve to reinforce the dominance of "traditional" heterosexuality, including women's subservient role in the household: they both aim to privilege men and punish individuals—both men and women—who do not embody heterosexual norms.

Structural Discrimination and the Health of Marginalized Populations

The negative effects of discrimination on individual health reach beyond interpersonal interactions to systemic and ecological factors (Collins et al. 2004; Dominguez 2008; Earnshaw et al. 2012; Mustillo et al. 2004). Structural discrimination encompasses multiple features of the social environment, including existing policies that

provide greater access to power and resources for certain groups and that differentially distribute power and resources. Consequently, laws and policies not directly related to health can potentially impact health (Taylor 2019). A robust body of literature has emerged that documents how LGB-specific policies (e.g., same-sex marriage bans, anti-LGB discrimination legislation) may harm sexual minority health (Everett et al. 2016; Hatzenbuehler 2009; Hatzenbuehler et al. 2010; Hatzenbuehler et al. 2011; Raifman et al. 2017).

Even more recently, Homan introduced the concept of structural sexism and documented its effect on women's health (Homan 2019). Homan (2019) argues that structural sexism encompasses political, cultural, and economic systems that favor men's access to power and resources; it does not have to be perceived for it to have a negative impact on an individual's health. Importantly, Homan's work on structural sexism also includes measures related to family planning policies, including abortion policy. Access to sexual and reproductive health services is essential to bodily autonomy and fundamentally reflects an individual's status within a population. Thus, access to family planning services can impact reproductive health via actual use of these services, but it is also an indicator of status in civil society more broadly. In fact, many states with the most restrictive family planning policies also have some of the worst health outcomes for women and their infants (Kawachi et al. 1999; Wallace et al. 2017).

Heteropatriarchy as a system of oppression that targets women, as well as sexual and gender minorities, may operate in ways that "spill over" and affect individuals who are not the intended target of these policies. For example, heterosexual women may be negatively impacted by environments that have more restrictive policies related to LGB rights, particularly if these policies are part of a broader political and social context that also seeks to regulate women and gender minorities. While previous studies have found that the impact of structural LGB stigma is concentrated among sexual minority persons (Hatzenbuehler 2017; Hatzenbuehler et al. 2014; Hatzenbuehler et al. 2020), one study found that school environments that were more LGB-exclusionary were associated with a younger age at first sex and lower rates of condom use among all adolescents (Philbin et al. 2021). None of these studies, however, have focused on birth outcomes, an outcome that may be uniquely sensitive to broader systems of stigma for several reasons.

The system of heteropatriarchy seeks to regulate both gender and sexuality, which likely disproportionately impacts women and sexual and gender minorities. However, while heteropatriarchy affects sexual minorities by constraining their access to power and resources, it uniquely targets people with the capacity for pregnancy by also constraining their access to family planning services. A handful of studies have identified a relationship between increased funding for family planning (including access to abortion) and improvements in adverse birth outcomes (Corman and Grossman 1985; Grossman and Jacobowitz 1981; Joyce 1987a, 1987b; Krieger et al. 2016; McFarlane and Meier 1998, 2001; Meier and McFarlane 1994). Other work has shown that women living in states with more reproductive rights have lower rates of both preterm and low birth weight births (Wallace et al. 2017). None of these studies examined the impact of such policies on the birth outcomes of sexual minorities.

To our knowledge, only one study has incorporated elements of both structural stigma and structural sexism (Charlton, Hatzenbuehler et al. 2019). In their study of sexually transmitted infections among adolescents and teen pregnancy risk, Charlton,

Hatzenbuehler et al. (2019) incorporated an indicator of the percentage of women without an abortion provider in their county into their scale, which also included measures of LGB-specific laws, percentage of same-sex couples at the state level, and public opinion polls toward LGB persons. The authors found that the abortion access indicator loaded well onto their measures of LGB stigma, and that their scale of structural stigma was associated with STI diagnosis among sexual minority adolescent females. Importantly, previous research has found that sexual minority women are more likely to report unintended pregnancy than their heterosexual counterparts (Charlton, Everett et al. 2019; Everett et al. 2017). Decreased access to family planning and abortion services, therefore, is likely to impact their birth outcomes at least as much as for their heterosexual peers.

Current Study

Our study explores the links between structural heteropatriarchy and birth outcomes for several reasons. Multiple studies have documented that birth outcomes respond to external stressors, including social policies and contexts. For example, changes in social policies, such as the repeal of Jim Crow laws and the passage of the Civil Rights Act, have been shown to improve maternal and child health among African Americans (Almond et al. 2006; Krieger et al. 2013). Other work has found a relationship between infant mortality and political context (e.g., Republican or Democratic administration). This includes decreases in infant mortality during Democratic administrations (Rodriguez et al. 2014); a negative relationship between political gender inequality (e.g., the percentage of state legislature seats occupied by women) and infant mortality (Homan 2017); and an association between women's political participation and economic autonomy, reproductive rights, and infant mortality (Kawachi et al. 1999; Koenen et al. 2006).

No research to date, however, has examined the impact of heteropatriarchy on birth outcomes (birth weight in kilograms, low birth weight, preterm birth). Building upon feminist theories of heteronormativity, we argue that social environments and policies related to the regulations of gender or sexuality create a synergistic system of structural heteropatriarchy. While previous research has considered these systems separately, such an approach may provide an incomplete picture of how a broader structural system that regulates gender and sexuality influences birth outcomes. We address this gap by using the National Longitudinal Study of Adolescent to Adult Health to create an index of heteropatriarchy that incorporates indicators of state-specific LGB policies, family planning policies, and women's political and economic position. We hypothesize that higher levels of heteropatriarchy will be associated with increased risk of adverse birth outcomes.

Data and Methods

Data and Sample

Data come from the National Longitudinal Study of Adolescent to Adult Health (Add Health). The initial Add Health sample was drawn from 80 high schools and 52

middle schools throughout the United States with unequal probabilities of selection in 1994 (Harris 2013; Harris and Urdy 2012). A subsample of students (*n* = 20,745) were asked to complete additional in-home interviews and were contacted for follow-up interviews in 2001–2002 (Wave III), 2007–2008 (Wave IV), and 2016–2018 (Wave V). Response rates were 80.3% for Wave IV and 69.3% for Wave V.

Respondents completed birth rosters in Waves IV and V, which covered dates of live births and birth outcome information, including birth weights and whether the infant was born preterm. We use births as the unit of analysis as this approach allows (1) each birth to be considered as a unique observation and (2) for the correct time ordering of various covariates owing to the longitudinal nature of the data (e.g., sexual identity of mother, educational attainment of mother, and parents' relationship status, all prior to pregnancy).

Our sample was restricted to live singleton births reported at Wave IV or V and to persons who identified as "female" at Wave I.[1] By Wave V, individuals with valid sample weights had reported a total of 13,217 live births. Additional births were excluded because they were not singleton (*n* = 268) or the maternal age was younger than the respondent's age at the first wave of data collection (*n* = 219). The sample sizes for birth weight, low birth weight, and preterm birth are 11,058, 11,320, and 11,414, respectively.

Measures

Birth Outcomes

Preterm birth was a dichotomous indicator of whether an infant was born earlier than 37 weeks' gestation (coded as "1") versus at 37 weeks' gestation or later (referent = 0). Both low birth weight and birth weight were derived from the same measure, which asked respondents to self-report the weight of their infant at birth. Low birth weight was constructed as a dichotomous indicator on the basis of whether an infant weighed less than 2,500 grams (coded as "1") or at least 2,500 grams (referent = 0). Birth weight was measured as a continuous measure of an infant's weight in kilograms. While these are self-reported indicators, previous studies have shown that respondents accurately report their infant's weight and gestational ages at birth (P. Dietz et al. 2014; Shenkin et al. 2017).

Heteropatriarchy Scale

The heteropatriarchy scale comprised 13 measures that represent three important dimensions of heteropatriarchy: (1) structural sexism, (2) family planning policy, and (3) LGB policy. All contextual measures were assessed at the wave indicated in Table 1 at the time of in-home interviews; GPS coordinates taken at the respondents' household were then linked to external contextual-level data sources (e.g., the U.S.

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[1] Gender identity was not assessed until Wave V of the survey.

Table 1 Indicators used to construct the heteropatriarchy scale, including overall Cronbach's alpha for the scale and individual alpha values for each indicator

Measure	Level	Wave	Alpha
Structural Sexism			
Ratio of men's/women's median income	County	4	.748
Ratio of men's/women's labor force participation	County	4	.762
Ratio of women's/men's unemployment rate	County	4	.763
% of votes cast for Republican president	County	1	.729
Conservative denomination adherents per capita	County	1	.713
Abortion Policy			
Public funding for abortion	State	1	.691
Abortion providers	County	1	.720
Mandatory waiting periods and informed consent	State	1	.732
Parental consent for abortion	State	1	.721
Lesbian, Gay, and Bisexual Policy			
Employment discrimination protection for sexual orientation	State	4	.721
Hate crime statute for sexual orientation	State	4	.723
Same-sex marriage/domestic partnership/civil union/reciprocal benefits relationship	State	4	.713
Same-sex adoption	State	4	.734
Test Scale Alpha			.745

Source: National Longitudinal Study of Adolescent to Adult Health.

Census, the Centers for Disease Control and Prevention, and election results). These contextual indicators came from ancillary contextual files available in the restricted Add Health data set.

Structural sexism included five indicators: (1) the census tract ratio of men's to women's median income; (2) the census tract ratio of men's to women's labor force participation; (3) the census tract ratio of women's to men's unemployment rate; (4) the county-level proportion of votes cast for the Republican presidential candidate; and (5) the county-level proportion of conservative denomination adherents (all religions combined).

The ratio for median income was constructed from two contextual indicators that measured median earnings for men and women aged 16 or older at Wave IV at the census tract level. Men's median income was divided by women's median income. Cases exceeding a value of five were omitted as they were extreme outliers (e.g., men making five times as much as women in that census tract). Finally, the measure was standardized so that a value greater than zero indicates that men were earning more than women. Similar measures were created to capture the ratio of men's to women's labor force participation and unemployment rate.

Rates for Republican presidential votes and conservative religious adherents per capita were not available at Wave IV, so Wave I values were used. The measure of votes cast for a Republican president was the proportion of voters who voted for the Republican candidate within a given county at Wave I; this value was standardized (Fowler et al. 2010). The proportion of Republican voters in an individual's neighborhood has been shown to impact the mental health of sexual minority adults (Everett 2014). Other research has demonstrated that conservative ideology is associated with

prejudice against LGB persons, in part because of the linkage between conservative ideology and the endorsement of traditional gender roles (Prusaczyk and Hodson 2018) and because hostile sexist attitudes predict voting for Republican presidential candidates (Bock et al. 2017; Schaffner et al. 2018). Conservative denomination adherents per capita was measured at the county level at Wave I. Conservative religious ideology has been associated with prejudice against LGB persons and has also been linked to the endorsement of traditional gender roles and sexist attitudes (Mikołajczak and Pietrzak 2014; Prusaczyk and Hodson 2018; Whitehead and Perry 2019).

We included four measures of family planning policy at Wave I (the only wave to collect such information), including three state-level indicators and one county-level indicator. Public funding for abortion captured whether a state provided funding for abortion in limited circumstances (e.g., rape, incest, or life endangerment; yes=1) or in all or most circumstances (referent=0). The measure of whether the respondent's county had an abortion provider was dichotomous (0=at least one provider, and 1= no provider). The presence of mandatory waiting periods and informed consent for abortion captured whether there was a state-level requirement that pregnant people receive lectures and state-prepared materials on fetal development, prenatal care, and adoption and observe a mandatory waiting period after such information is received. States with no informed consent laws were coded as "0"; states with informed consent laws and an unenforced mandatory waiting period were coded as "1"; and states with informed consent laws and enforced waiting periods were coded as "2." Parental consent for abortion was measured at the state level and captured whether a state had no parental consent laws (coded as "0"), unenforced consent laws (coded as "1"), or enforced consent laws (coded as "2").

We used four Add Health indicators of LGB policy in a given context derived from data from the Human Rights Council; these data were released in 2019 as part of an ancillary study of LGB contexts. Four dichotomous indicators measured whether a state had various sexual orientation–related policy protections at Wave 4. For unemployment discrimination, states were coded as not having protections (1) if they did not prohibit employment discrimination based on sexual orientation in private or public employment (referent=0). A state was coded as "1" if it did not have hate crime statutory provisions based on sexual orientation. States that did not allow same-sex marriage, domestic partnerships, civil unions, or reciprocal beneficiary relationships were coded as "1." Finally, states that did not allow same-sex joint adoption or second-partner adoption were coded as "1."

A summary of the heteropatriarchy measure is provided in Table 1. The higher the value on the various scales, the higher the level of heteropatriarchy present in the respondent's state and county environment. Table 1 also provides details about each measure—the level of analysis that it represents, the wave of data from which it was derived, and its individual alpha value. A test of internal consistency of all 13 measures yielded a Cronbach's alpha of .75.

Control Variables

Maternal age was derived from each respondent's date of birth and the month and year of their infant's birth. Maternal age ranged from 18 to 41.

Sexual identity was assessed in Waves III, IV, and V. Respondents were asked, "Please choose the description that best fits how you think about yourself: 100%

heterosexual (straight); mostly heterosexual (straight), but somewhat attracted to people of your own sex; bisexual, that is, attracted to men and women equally; mostly homosexual (gay), but somewhat attracted to people of the opposite sex; 100% homosexual (gay); not sexually attracted to either males or females; or missing." Respondents who identified as lesbian or bisexual were collapsed into one category owing to small sample sizes. We used the sexual identity measure prior to the pregnancy.

Education level prior to pregnancy was measured using data from all waves. Respondents were asked about the highest degree they had received at each wave, and in Wave IV, the year they received that degree. From these data, we calculated the age at which they finished their degree and whether degree completion occurred prior to or after each pregnancy. Educational achievement was measured as a categorical variable that captured whether, prior to the pregnancy, the respondent had (1) less than a high school degree, (2) a high school degree or equivalent, (3) some college, or (4) a bachelor's degree or more.

Relationship status was assessed by asking the following for each pregnancy reported at Wave IV and for all live births reported at Wave V: "Were you married to [pregnancy partner] at the time of the pregnancy/birth?" If respondents answered no, they were then asked, "Were you and [pregnancy partner] living together at the time of pregnancy/birth?" If respondents answered no, they were then asked, "Which of the following best describes your relationship with [pregnancy partner] at the time of pregnancy/birth: we did not see or talk to each other; we hardly ever saw or talked to each other; we were just friends; we were involved in an on-again, off-again relationship; we were romantically involved on a steady basis." From these responses, we created a categorical variable that measured relationship to pregnancy partner as married (referent), cohabiting, dating, just friends/stranger, and unknown.

Race and ethnicity were measured categorically as non-Hispanic White (referent), non-Hispanic Black, Latina, and other race or ethnicity.

Poverty was measured at the census tract level at Wave IV and reflected the proportion of individuals living in poverty in a given state at that time. This measure was converted to quintiles, with the first quintile (the lowest level of poverty) serving as the reference category.

Region was a four-category indicator of U.S. geographic region. This was included as a control since previous research has shown that some regions have relatively high levels of structural sexism, limited protections of LGB individuals, and less family planning access. To protect the identity of respondents, the actual regions to which the values correspond were unknown.

Analytic Strategy

First, we present descriptive statistics for the total sample. To examine the relationship between heteropatriarchy and birth outcomes, we conducted a series of analyses using ordinary least-squares regression for our analysis of birth weight and logistic regression for our analyses of preterm birth and low birth weight using clinical cutoffs. We present the bivariate relationship between heteropatriarchy and birth outcomes (model 1), followed by a multivariate model that includes all of our controls

(model 2). All models adjust for population weights and cluster on respondent's ID to account for shared variance across multiple births to the same person.

Results

Table 2 presents descriptive statistics for the analytic sample. The mean score for the heteropatriarchy scale was 6.51 in our sample. While the mean scores for men's to women's labor force participation and unemployment were 0, indicating roughly equal rates, on average, births in this sample occurred in counties with more Republican voters and conservative religious adherents. The majority of births were to respondents who lived in states with employment discrimination protections (55%) and hate crime protections (59%) based on sexual orientation. However, it was much less common for births to be reported by respondents who lived in states with same-sex marriage protections (14%) and legal same-sex adoption (20%). The majority of births were reported in states with limited abortion protections at Wave I: 70% of respondents lived in states with limited public funding for abortion and 41% lived in a county with no abortion provider. Additionally, the need for both informed consent and parental consent was common in the states in which respondents lived. Nearly 46% of births were reported by individuals who lived in a state where informed consent and mandatory waiting periods were required, while 64% lived in states where parental consent was enforced.

The majority of births in the sample were reported by respondents who identified as 100% heterosexual (83%), White (68%), married at the time of birth (55%), and high school graduates (47%). About 23% were cohabiting and 16% were dating their partners at the time of pregnancy. The average age at pregnancy was 26. In our sample, 12% of births were preterm, 7% of births were low birth weight, and the average infant weight was 3.31 kg.

Table 3 presents multivariate models for the impact of heteropatriarchy on birth outcomes. In panel A, which gives the results for preterm birth, model 1 shows that as the level of heteropatriarchy increases, so does the risk for preterm birth (OR $=1.04$, $p<.001$). This relationship persists after the inclusion of controls in model 2. Figure 1 provides the predicted probabilities of a preterm birth relative to the level of heteropatriarchy in a given context with all other predictors held at their means. For births reported by women living in contexts with a score of 1 on our heteropatriarchy scale, the probability of reporting a preterm birth is .08, less than half of the probability of reporting a preterm birth among women living in states at the high end of the heteropatriarchy scale.

As shown in panel C of Table 3, as the level of heteropatriarchy increases, birth weight decreases ($B=-0.01$, $p<.001$); this finding persists in model 2 ($B=-0.01$, $p<.05$). Figure 2 presents the predicted birth weight as a function of heteropatriarchy level. Results show that the predicted birth weight for an infant born under the lowest level of heteropatriarchy is 3.38 kg (95% CI=3.31–3.44), while it is 3.22 kg at the highest level of heteropatriarchy (3.14–3.30). We did not find a significant relationship between heteropatriarchy and low birth weight in our bivariate or multivariate analyses (panel B).

Table 2 Descriptive statistics for the analytic sample

Measure	Mean/Ratio/%
Heteropatriarchy Scale (mean; range, 0–13)	6.51 (0.32)
Structural sexism	
Ratio of men's/women's median income	0.04 (0.01)
Ratio of men's/women's labor force participation	0.00 (0.01)
Ratio of women's/men's unemployment rate	0.00 (0.02)
% of votes cast for Republican president	0.10 (0.09)
Conservative denomination adherents per capita	0.19 (0.11)
Lesbian, gay, and bisexual policy (%)	
Employment discrimination protection for sexual orientation	55.14
Hate crime statute for sexual orientation	58.52
Same-sex marriage/domestic partnership/civil union/reciprocal benefits relationship	14.08
Same-sex adoption legal	19.91
Abortion policy (%)	
Limited public funding for abortion	69.90
No abortion provider	40.56
Informed consent	
None	36.02
Informed consent but no mandatory waiting period	18.35
Informed consent and mandatory waiting period	45.63
Parental consent	
None	18.84
Not enforced	17.27
Enforced	63.89
Birth Outcome	
Preterm birth (%)	11.95
Low birth weight (%)	7.23
Birth weight (kg)	3.31 (0.01)
Additional Covariates	
Maternal age (years)	26.20 (0.20)
Sexual identity (closest) (%)	
100% heterosexual	82.86
Mostly heterosexual	14.09
Mostly or 100% gay/bisexual	2.53
Not reported	0.51
Education (prior to birth) (%)	
Less than high school	13.25
High school (or equivalent)	47.18
Some college	16.10
Bachelor's degree or more	23.47
Relationship status (%)	
Married	54.85
Cohabiting	22.76
Dating	16.27
Just friends/stranger	6.03
Unknown	0.09
Race and ethnicity (%)	
Non-Hispanic White	68.14
Non-Hispanic Black	17.21
Latina	10.94
Other	3.71

Table 2 (continued)

Measure	Mean/Ratio/%
Poverty (%)	
First quintile	28.57
Second quintile	21.51
Third quintile	23.82
Fourth quintile	26.10
Region (%)	
1	14.70
2	34.66
3	40.44
4	10.20

Notes: Analytic sample $N=11,414$. Figures are means, ratios, or percentages, and standard errors are shown in parentheses.

Source: National Longitudinal Study of Adolescent to Adult Health.

Supplementary Analyses

We conducted a series of sensitivity analyses. First, we explored a series of interactions between heteropatriarchy and sexual identity, as well as between heteropatriarchy and race and ethnicity. We did not find significant interactions between our measure of structural heteropatriarchy and these characteristics. Because sexual orientation is multidimensional, we also conducted a series of sensitivity analyses that included indicators of same-sex sexual attraction and same-sex sexual behaviors, as well as interactions between our heteropatriarchy scale and these indicators. These measures were not significant in our models and neither were the interactions. Taken together, these results suggest that the impact of heteropatriarchy on birth outcomes does not significantly vary across respondents.

Second, because it is possible that respondents moved between the Wave 1 and Wave IV surveys from which our contextual measures were derived, we conducted a series of analyses that restrict our sample to people who did not move more than 50 miles away from their georeferenced home location in Wave I (see Table A1 in the online appendix). These results continue to show a significant relationship between heteropatriarchy and adverse birth outcomes. Among respondents who did not move, a one-unit increase in structural heteropatriarchy was associated with an increased risk of a preterm birth (OR$=1.04$, $p<.05$) and a lower birth weight ($B=-0.01$, $p<.10$). This suggests that moving does not alter the impact of being exposed to these various forms of heteropatriarchy.

Finally, we conducted analyses that were restricted to births that occurred prior to 2010 (see Table A2). Although we incorporated births reported in Wave V, this wave's contextual measures were not available at the time of publication. To account for differences in exposure to heteropatriarchy between Waves IV and V, in conducting these analyses we excluded births that occurred after the Wave IV interview. Our results are also robust to this specification. A one-unit increase in structural heteropatriarchy was associated with an increase in the risk of a preterm birth (OR$=1.04$, $p<.05$) and a lower birth weight ($B=-0.01$, $p<.01$).

Table 3 Multivariate results (odds ratios and coefficients) assessing associations between level of heteropatriarchy and preterm birth, low birth weight, and birth weight (kg)

Measure	A. Preterm Birth		B. Low Birth Weight		C. Birth Weight	
	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2
Heteropatriarchy	1.04***	1.04*	1.02	1.01	−0.01***	−0.01*
	(1.02–1.06)	(1.01–1.07)	(1.00–1.05)	(0.98–1.05)	(−0.01 to 0.00)	(−0.01 to −0.00)
Maternal Age		0.99		0.99		0.00
		(0.98–1.01)		(0.97–1.01)		(−0.00 to 0.01)
Sexual Identity (ref. = 100% heterosexual)						
Mostly heterosexual		0.87		0.91		0.02
		(0.66–1.15)		(0.63–1.30)		(−0.03 to 0.07)
Mostly or 100% gay/bisexual		0.59*		0.89		0.05
		(0.36–0.99)		(0.51–1.54)		(−0.04 to 0.14)
Not reported		0.41		0.33*		−0.09
		(0.13–1.31)		(0.11–1.00)		(−0.29 to 0.11)
Education (ref. = less than high school)						
High school		1.11		1.06		0.03
		(0.84–1.47)		(0.80–1.40)		(−0.02 to 0.08)
Some college		1.04		0.98		0.05
		(0.74–1.46)		(0.68–1.43)		(−0.01 to 0.11)
Bachelor's degree or more		0.73		0.88		0.07*
		(0.50–1.06)		(0.58–1.33)		(0.00–0.13)
Relationship Status (ref. = married)						
Cohabiting		1.06		1.47**		−0.08**
		(0.86–1.30)		(1.13–1.91)		(−0.12 to −0.03)
Dating		0.65**		0.95		−0.02
		(0.50–0.85)		(0.70–1.30)		(−0.07 to 0.03)
Just friends/stranger		0.87		1.21		−0.06†
		(0.62–1.22)		(0.82–1.78)		(−0.13 to 0.00)
Unknown		3.07		—		0.03
		(0.44–21.56)		—		(−0.14 to 0.21)

Table 3 (continued)

Measure	A. Preterm Birth		B. Low Birth Weight		C. Birth Weight	
	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2
Race (ref.=non-Hispanic White)						
Non-Hispanic Black		1.05		1.90***		−0.17***
		(0.83–1.33)		(1.47–2.46)		(−0.23 to −0.12)
Hispanic		0.91		1.06		−0.07*
		(0.68–1.21)		(0.76–1.47)		(−0.12 to −0.01)
Other		0.94		0.98		−0.09†
		(0.59–1.50)		(0.56–1.72)		(−0.19 to 0.01)
Poverty (ref.=first quintile)						
Second quintile		0.78†		0.74†		0.02
		(0.61–1.00)		(0.53–1.02)		(−0.04 to 0.07)
Third quintile		1.01		0.88		−0.04
		(0.79–1.30)		(0.66–1.19)		(−0.09 to 0.02)
Fourth quintile		0.99		0.91		0.00
		(0.76–1.29)		(0.67–1.23)		(−0.05 to 0.06)
Region (ref.=region 1)						
2		0.83		0.73†		0.01
		(0.61–1.13)		(0.50–1.06)		(−0.05 to 0.08)
3		0.86		0.84		−0.01
		(0.63–1.19)		(0.57–1.22)		(−0.09 to 0.06)
4		0.87		0.69†		−0.01
		(0.62–1.21)		(0.47–1.02)		(−0.08 to 0.07)
Constant	0.10***	0.16***	0.07***	0.11***	3.36***	3.32***
	(0.09–0.12)	(0.08–0.31)	(0.06–0.09)	(0.05–0.24)	(3.33–3.40)	(3.19–3.45)
No. of Observations	11,410	11,410	11,319	11,311	11,236	11,236

Notes: Robust 95% confidence intervals are shown in parentheses.

Source: National Longitudinal Study of Adolescent to Adult Health.

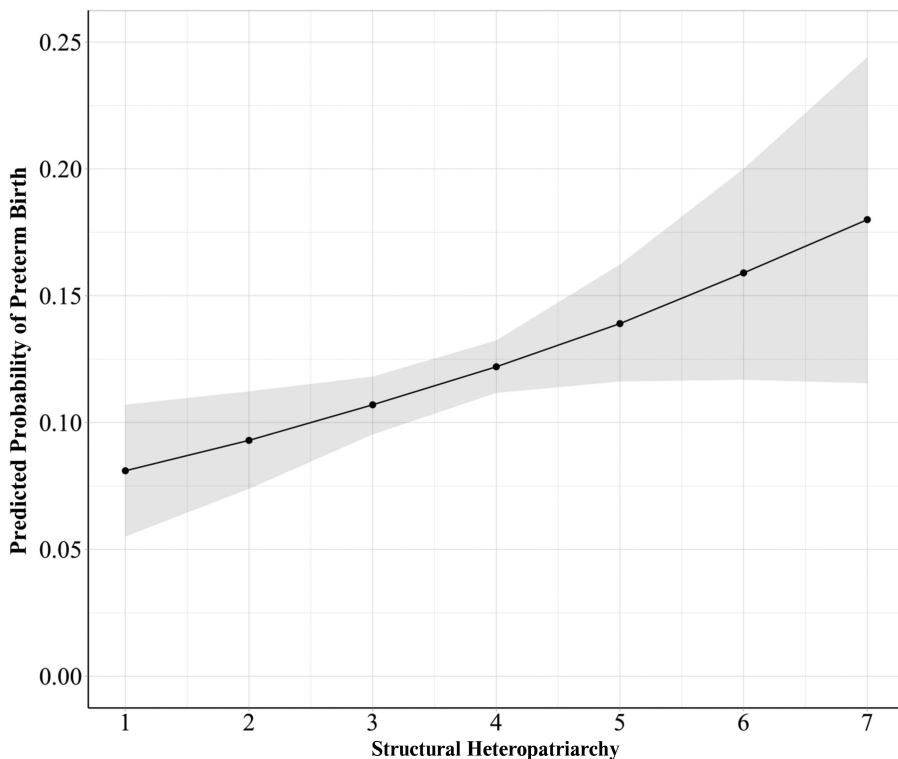†$p<.10$; *$p<.05$; **$p<.01$; ***$p<.001$

Fig. 1 Predicted probability of preterm birth as a function of the level of heteropatriarchy. The shaded area represents the 95% confidence interval.

Discussion

The deployment of both gender and sexuality in maintaining systematic gender oppression has a long and well-established history in feminist theory (Butler 1991; Rich 1980; Rubin 1975), and contemporary studies have reaffirmed this in modern contexts (Westbrook and Schilt 2014). These theories argue that it is impossible to comprehensively understand the harmful effects of discrimination based on sex and gender and discrimination based on sexual orientation as separate systems. Our study is the first to synergistically incorporate measures of structural sexism, family planning policy, and structural LGB-based discrimination and examine their impact on birth outcomes.

Our results have several implications for demographic research. First, we found a high level of correlation between our indicators of structural sexism, family planning policies, and LGB policy measures. These findings further bolster feminist scholars' claims regarding the interconnected nature of heterosexuality and gender oppression (Butler 1991; MacKinnon 1982; Rubin 1998; Schilt and Westbrook 2009). Whereas previous work has illuminated how sexist and heterosexist norms are both implicated in the reproduction of inequality at the interpersonal level (Schilt and Westbrook 2009), our findings document the interdependence of these systemic inequalities at the institutional level. As demographic researchers increasingly focus on the impact of
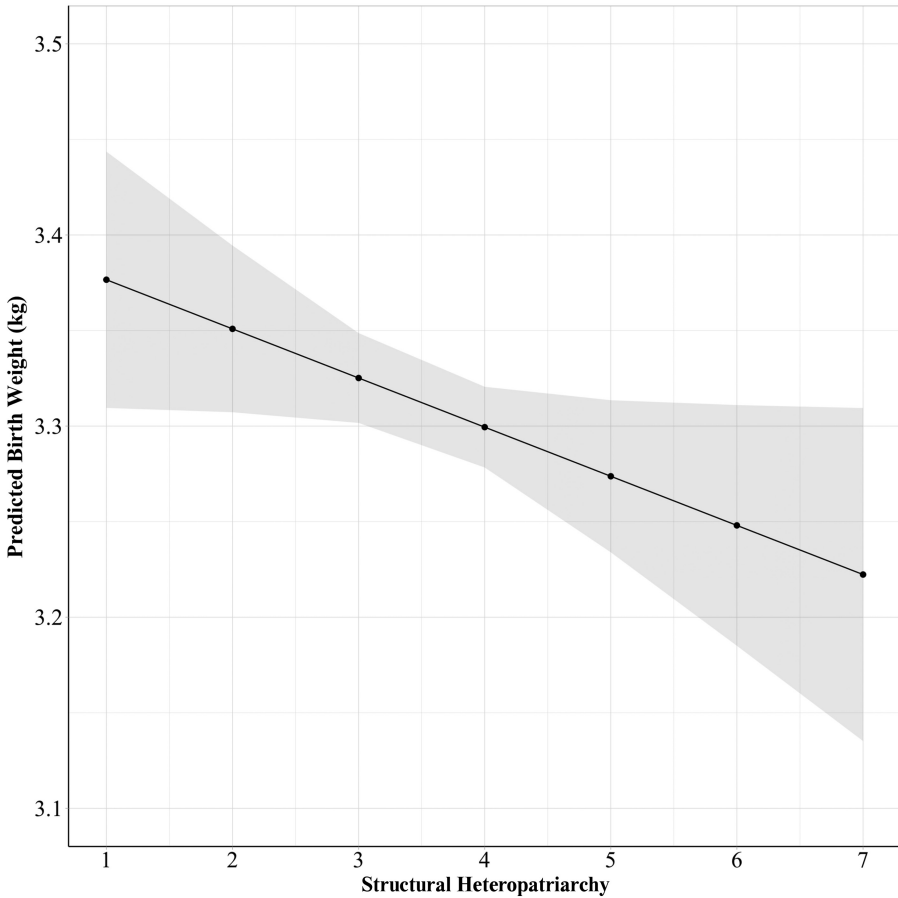
Fig. 2 Predicted birth weight as a function of the level of heteropatriarchy. The shaded area represents the 95% confidence interval.

structural forms of discrimination on individual health, our results highlight the need to think of different forms of identity-based stigma as not necessarily independent but as a part of a larger system of discrimination that privileges White, cisgender men.

Second, we found evidence that higher levels of heteropatriarchy were associated with increased risk of preterm birth and lower birth weight. Results for the clinical cutoff of low birth weight were not significant. We do not find this surprising given new evidence that birth weight has been declining in the United States owing to increased reliance on obstetric interventions (e.g., cesarean deliveries) (MacDorman et al. 2010). We did not find moderating effects by sexual identity or race and ethnicity, suggesting that the harmful effects of heteropatriarchy on pregnancies exist regardless of the mother's race, ethnicity, or sexual identity. Our study expands upon research reporting that exposure to discrimination and stress negatively affect birth outcomes (Almond et al. 2006; Christian 2012; Gemmill et al. 2019). Pregnancy is a unique state during which an individual's intergenerational stress exposure has implications for health. Moreover, a pregnant person need not "feel" their heteropatriarchal environment as potentially detrimental for themselves or their pregnancy for that preg-

nancy to be negatively impacted. Indeed, systemic heteropatriarchy can limit access to reproductive health care or shape the kind of care offered; it can limit access to socioeconomic opportunities; and it can shape the norms and expectations around reproduction, all of which might contribute to adverse birth outcomes.

Importantly, our measure of heteropatriarchy also incorporates indicators of family planning access. In recent years, the number of policies passed to restrict abortion access has grown: between 2010 and 2018, states passed 424 policies that restrict access to abortion services (Nash et al. 2018). These restrictive policies likely directly and indirectly influence birth outcomes in multiple ways. Family planning policies allow individuals to decide whether and when to have children. People often avoid childbearing for reasons that are related to birth outcomes, including financial insecurity, issues related to mental health, and abusive romantic partners (Foster et al. 2012). People also terminate pregnancies because of health issues directly related to the pregnancy itself; pregnant people who are unable to access family planning services, therefore, may be forced to have births that are more likely to end in adverse birth outcomes. Our results also suggest that abortion policy is part of a larger system of sexism and heteropatriarchy that impacts birth outcomes (Hodson and MacInnis 2017; Nossiff 2007).

These structural barriers to self-determination that people face as a result of restrictive abortion policies have much in common with policies that restrict the ability of sexual minorities to achieve full citizenship in the United States. Although the country recently legalized same-sex marriage, and in June 2020 made it illegal to fire someone based on their sexual orientation or gender identity, in many states it is still legal to deny adoption rights or restrict access to housing based on sexual orientation. Together, these seemingly separate sets of regulations undergird larger heteropatriarchal social arrangements to negatively impact birth outcomes. Demographers should continue to incorporate indicators of abortion policies on reproductive outcomes, but may want to consider the implications of these policies as indicators of structural stigma that may impact other health- and socioeconomic-related outcomes.

Limitations and Conclusions

Our study has several limitations that should be noted. First, our measures of family planning, percentage of conservative religious adherents, and percentage of Republican voters were derived from Wave I of the survey, which may not reflect the social environment during the pregnancy. We conducted a series of sensitivity analyses that restrict our sample to respondents who did not move more than 50 miles from their home address measured at Wave I, and the results were robust to these specifications, which suggests two things. Across the country, access to abortion has become *more* restrictive, thus family planning policies at Wave I likely underestimate the restrictions individuals may be exposed to later in life. It is also unlikely that substantial changes in the political or religious landscape have occurred in a given community.

A second limitation is that, even if respondents did move from an extremely conservative environment to an extremely liberal one, the fact that our results hold in *both* samples (nonmovers and movers) suggests that early developmental environments have potentially important long-term consequences for women's reproductive health. Other research has shown an important relationship between

early life context (childhood and adolescence) and health in later life (Brooks-Gunn et al. 1993; Elder 1998; Hayward and Gorman 2004; Wickrama and Noh 2010). Our study is also limited by the fact that we rely on self-reported data, however, research has shown that women provide accurate reports of their infants' birth weights and gestational ages (P. Dietz et al. 2014; P. M. Dietz et al. 2014).

Finally, we did not have adequate power or measurement of gender identity among our respondents. Add Health asks about gender identity only in Wave V and, because of small sample sizes, respondents who did not answer the survey item and those who selected an identity other than "male" or "female" were combined. In our sample, only 17 births were to respondents who either did not answer the survey item or selected the "other" category. No respondents selected a "male" gender identity at Wave V. While we did not find that the relationship between our structural heteropatriarchy and birth outcomes was modified by sexual orientation, there may be differential effects by gender identity that warrant future investigation.

Future research should consider additional ways to refine and expand the measurement of structural heteropatriarchy. While the regulation of gender expression through policy related to transgender rights is clearly relevant to the construct of heteropatriarchy, Add Health does not currently include any policy-level data on gender identity or transgender rights. Data from the Movement Advancement Project, which provides state-specific data on LGB and gender identity policies, reveal a high correlation between states with restrictive policies based on sexual identity and policies based on gender identity (Movement Advancement Project 2021). We hypothesize that policies that limit or expand rights for transgender persons would improve the strength of our heteropatriarchy scale or identify differences that might redefine it as a cis-heteropatriarchy scale. Moreover, this line of research would also benefit from the investigation of the links between structural racism and heteropatriarchy ("White supremacist heteropatriarchy") and how these systems of inequality jointly shape population health. Recent intersectional scholarship points to connections between heteropatriarchy, colonialism, and White supremacy (Arvin et al. 2013; Hooks 2000; Valdes 1996).

Despite these limitations, our study is the first to combine indicators of structural gender- and sexual orientation–based discrimination to create an indicator of heteropatriarchy. We show that not only do these indicators overlap to create a parsimonious construct, but that exposure to heteropatriarchy—even if only in early life—has negative consequences for birth outcomes. This study builds on and extends conversations around the role of social environments by encouraging researchers to move beyond examining gender and sexuality as independent systems of domination, but rather to explore them as co-constitutive and reinforcing systems that can be conceptualized as a single heteropatriarchal structure with important implications for population health disparities. ∎

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