

Structural Sexism and Health in the United States

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Dissertation submitted in partial fulfillment of  
the requirements for the degree of Doctor  
of Philosophy in the Department of  
Sociology in the Graduate School  
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2018

ABSTRACT

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## Abstract

In this dissertation, I seek to begin building a new line of health inequality research that parallels the emerging structural racism literature by developing theory and measurement for the new concept of *structural sexism* and examining its relationship to health. Consistent with contemporary theories of gender as a multilevel social system, I conceptualize and measure structural sexism as *systematic gender inequality in power and resources* at the macro-level (U.S. state), meso-level (marital dyad), and micro-level (individual). Through a series of quantitative analyses, I examine how various measures of structural sexism affect the health of men, women, and infants in the U.S.

Chapter 2 focuses exclusively on the macro-level and drills down on a single indicator of structural sexism: political gender inequality. Using data compiled from several administrative sources across a twenty-two-year period, I find that women's political representation in U.S. state legislatures is as strong — or stronger — a predictor of state infant mortality rates than even the state's poverty level. This relationship holds both between states with varied levels of political gender inequality in a given year, and within the majority of states as women's representation changes over time. To quantify the population-level costs, I estimate that women's continued underrepresentation in legislative office was associated with nearly 3,500 excess infant deaths per year in the U.S. compared to what would be expected under conditions of gender parity.

Chapter 3 lays out a more comprehensive, multilevel framework for structural sexism and examines how it shapes the health of women and men at midlife. I measure macro-level structural sexism at the U.S. state-level using indicators of inequality in political, economic, cultural and reproductive domains. Using restricted geo-coded data from the NLSY79, individuals are located within states to capture their exposure to structural sexism. This chapter also incorporates individual- and spousal-level data from the NLSY79 in order to measure exposure to structural sexism at the meso- and micro-levels. Results show that among women exposure to more sexism at the macro- and meso-levels is associated with more chronic conditions, worse self-rated health, and worse physical functioning. Among men, macro-level structural sexism is also associated with worse health. However, at the meso-level greater structural sexism is associated with better health among men. At the micro-level, internalized sexism is not related to health among either women or men. These results highlight the importance of a multilevel approach.

As a whole, this work demonstrates that structural sexism takes a substantial toll on population health in the U.S. The findings also illustrate that effects of structural sexism on health depend both on an individual's position within the gender system and the level at which sexism exposures are measured. Further research is needed to extend our understanding of how structural sexism shapes health across the life course.

## **Dedication**

To my #1 fans: Taylor, Wesley, Graham and Rusty. Thanks for cheering me on.

# Contents

Abstract.....	iv
List of Tables .....	ix
List of Figures .....	xi
Acknowledgements .....	xii
1. Introduction .....	1
2. Political Gender Inequality and Infant Mortality in the United States, 1990-2012.....	4
2.1. Background .....	5
2.2. Methods .....	11
2.2.1. Sample.....	11
2.2.2. Measures .....	12
2.2.3. Analytic Methods .....	14
2.3. Results .....	17
2.4. Discussion.....	27
3. This Gender System Makes Me Sick! Structural Sexism and Health at Midlife .....	33
3.1. Gender Discrimination and Health.....	35
3.2. Gender Differences (Gaps) in Health Outcomes .....	38
3.3. Toward a Structural Sexism and Health Approach .....	42
3.3.1. Macro-Structural Sexism .....	46
3.3.2. Meso-Structural Sexism.....	47
3.3.3. Micro-Structural Sexism .....	50
3.4. Methods .....	52

3.4.1. Sample .....	52
3.4.2. Measures .....	54
3.4.3. Analytic Methods .....	61
3.5. Results .....	63
3.5.1. A Descriptive Picture of Multilevel Structural Sexism in the U.S. ....	63
3.5.2. The Relationship Between Structural Sexism and Health at Midlife.....	68
3.6. Discussion and Conclusions .....	74
4. Conclusion.....	80
Appendix A.....	82
Appendix B .....	92
References.....	104
Biography .....	120

## List of Tables

Table 1. U.S. State-Level Descriptive Statistics, 1990-2012.....	17
Table 2. Zero-Order Correlations Between Key Variables.....	19
Table 3. Results of Longitudinal Regression Models Predicting U.S. State Infant Mortality Rates, 1990-2012.....	25
Table 4. Health Problems by Gender at Age 40 and 50 among all Individuals and Among Married Individuals.....	55
Table 5. Measures of Macro-Structural Sexism.....	57
Table 6. Individual Level Descriptive Statistics (For Person-Years in Full Sample and Married Subsample).....	61
Table 7. State-Level Descriptive Statistics, 1998-2012 (Among Person-Years in Full Sample).....	64
Table 8. U. S. States with the Highest and Lowest Levels of Sexism (Based on State Averages 1998-2012).....	66
Table 9. Meso-Level Descriptive Statistics (For Married Person-Years, n=4336).....	67
Table 10. Micro-Level Sexism Index Scores by Gender (For Person-Years in Full Sample and Married Subsample).....	68
Table 11. Coefficient Estimates For Structural Sexism From Random Effects Models Predicting Women's Physical Health Outcomes.....	70
Table 12. Coefficient Estimates for Structural Sexism Exposure from Random Effects Models Predicting Men's Physical Health Outcomes.....	71
Table 13. Results of Longitudinal Regression Model Predicting U.S. State Infant Mortality Rates Including Political Party Variables, 1990-2012.....	85
Table 14. Results of Longitudinal Regression Models Predicting U.S. State Infant Mortality Rates with Alternative Lag Periods, 1990-2012.....	87
Table 15. Variance Inflation Factors for All Variables in Full Model (R4).....	88

Table 16. Alternative Specifications of Fixed-Effects Regression Models Predicting Infant Mortality Rates, United States, 1990-2012 .....	89
Table 17. Alternative Specifications of Random-Effects Regression Models Predicting Infant Mortality Rates, United States, 1990-2012.....	91
Table 18. Comparison of Macro-Sexism Results Using Alternative Specifications of Index Among Married Women.....	93
Table 19. Comparison of Macro-Sexism Results Using Alternative Specifications of Index Among Married Men.....	94
Table 20. Leave-One-Out Replication of Macro-Sexism Results Among Married Women .....	95
Table 21. Leave-One-Out Replication of Macro-Sexism Results Among Married Men...	96
Table 22. Individual Macro-Sexism Indicators Among Married Women .....	97
Table 23. Individual Macro-Sexism Indicators Among Married Men .....	98
Table 24. Comparison of Meso-Sexism Results Using Alternative Specifications of Index Among Married Women.....	101
Table 25. Comparison of Meso-Sexism Results Using Alternative Specifications of Index Among Married Men.....	102

## List of Figures

Figure 1 IMR by Percentage Women in Legislature, U.S. States, 1990-2012 (n=600) .....	20
Figure 2. Within-State Relationships Between IMR and Women in Legislature, 1990-2012 .....	21
Figure 3 Hypothetical Relationships Between Structural Sexism and Health Among Women and Men.....	41
Figure 4 Conceptual Model of Structural Sexism.....	45
Figure 5 Macro-Structural Sexism Index, State Averages 1998-2012.....	65
Figure 6 Predicted Values of Health Outcomes Given Structural Sexism Exposure .....	73
Figure 7. IMR by Selected Controls, U.S. States, 1990-2012 (n=600) .....	83

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# 1. Introduction

In this dissertation, I seek to begin building a new line of health inequality research that parallels the emerging structural racism literature by developing theory and measurement for the new concept of *structural sexism* and examining its relationship to health. The nascent structural racism literature grows out of the recognition that measures of perceived discrimination—while illuminating an important piece of the puzzle — stop short of capturing the full effects of racism on health because racism is not exclusively an interpersonal level phenomenon and because a large portion of racial discrimination is unperceived (Bonilla-Silva 1997; Gee and Ford 2011). Similarly, gender theorists have long argued that gender is not primarily an attribute of individual bodies, but rather a multilevel social system (Ridgeway and Correll 2004; Risman 2004). Much gender inequality is unperceived and/or recast as natural results of inherent gender differences. Despite these similarities between race and gender systems, the application of the structural discrimination and health approach to the study of gender has been quite limited (Krieger 2014). Therefore, I advance a structural sexism and health perspective that is grounded in contemporary understandings of gender as a social structure and shifts the focus from comparisons between men and women on health outcomes, to within-gender comparisons across levels of exposure to discriminatory gender systems. Toward that end, the major contribution of this dissertation work is to

develop measures of structural sexism at the macro-, meso-, and micro-levels, and to examine how they affect the health of men, women, and infants.

Chapter 2 focuses exclusively on the macro-level and drills down on a single indicator of structural sexism: political gender inequality. In this chapter, I use data compiled from several administrative sources across a twenty-two-year period, to examine the relationship between women's political representation in U.S. state legislatures and state infant mortality rates (IMRs). I investigate this relationship both between states with varied levels of political gender inequality in a given year, and within states as women's representation changes over time. To quantify the population-level costs of women's continued underrepresentation, I compare the predicted IMR at the mean observed level of women's legislative representation in 2012 to the predicted IMR under hypothetical conditions of gender equality, and then calculate the number of excess infant deaths attributable to political gender inequality each year in the U.S.

Chapter 3 lays out a more comprehensive, multilevel framework for structural sexism and examines how it shapes the health of women and men at midlife. I measure macro-level structural sexism at the U.S. state-level using indicators of inequality in political, economic, cultural and reproductive domains. Using restricted geo-coded data from the NLSY79, individuals are located within states to capture their exposure to structural sexism. This chapter also incorporates individual- and spousal-level data from the NLSY79 in order to measure exposure to structural sexism at the meso- and

micro-levels. I first describe the extent and patterns of structural sexism exposure at the macro-, meso-, and micro-levels. I then examine how the structural sexism exposures at each level are related to chronic conditions, self-rated health, and physical functioning among both women and men in midlife. Finally, I consider whether the patterns of health effects support a conflict perspective (in which the dominant group benefits while the subordinate group suffers), a theory of universally harmful inequality (in which everyone suffers, although perhaps to varying degrees), or neither. Interestingly, the macro-level results support a different perspective than the meso-level results, highlighting the importance of a multilevel approach. Taken together, the results demonstrate the utility of a structural sexism approach to gender inequality and health.

## **2. Political Gender Inequality and Infant Mortality in the United States, 1990-2012<sup>1</sup>**

The United States' infant mortality rate (IMR) is among the highest in the developed world. According to the CDC, the U.S. had an IMR of 6.1 per 1,000 live births in 2010, ranking 26<sup>th</sup> out of the 29 OECD countries (MacDorman et al. 2014). The reasons for the United States' poor performance on this key indicator of population health are not fully understood, but improvements in the IMR have been a national priority for decades. In 2012, the U.S. Department of Health and Human Services began developing the first national strategy to address infant mortality (Lu and Johnson 2014). The proposed strategy recognizes the importance of addressing the social determinants of maternal and infant outcomes; however, gender inequality (or women's status in society) is notably absent from the list of social determinants considered. This omission likely stems from the fact that research on the health implications of women's status has focused almost exclusively on the less-developed world. The few studies that have examined women's status in the U.S. have found harmful effects of gender inequality in social and economic arrangements on the health of men, women, and children (Chen et al. 2005; Kawachi et al. 1999; Koenen, Lincoln, and Appleton 2006; Yllö 1984) but have not identified a consistent link between gender inequality in the political domain and infant mortality. Furthermore, all of these studies relied on cross-sectional data. The

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present study is the first to document the relationship between gender inequality in state legislatures and state infant mortality rates in the U.S. over time.

## **2.1. Background**

Gender equity has long been recognized as a crucial factor in health, development, and poverty reduction in the developing world (World Bank 2003). There also is general consensus that improving women's social and economic position benefits their children (Heaton 2015; World Bank 2003). Several aspects of child well-being-- including nutrition, immunization rates, child mortality under age five, and infant mortality-- are associated with various measures of women's status (Heaton 2015; Pratley 2016; Swiss, Fallon, and Burgos 2012). In particular, cross-national comparative work focused on the determinants of IMRs in less-developed countries has repeatedly linked higher women's literacy rates and educational attainment to reductions in infant mortality (Boehmer and Williamson 1996; Caldwell 1986; Frey and Field 2000; Shandra et al. 2004; Shen and Williamson 2001). Some of these studies also identified a negative relationship between women's political empowerment (as measured by parliamentary seats and suffrage rights) and IMRs (Boehmer and Williamson 1996; Quamruzzaman and Lange 2016; Swiss et al. 2012). Finally, Boehmer and Williamson (1996) also found that IMRs in less developed countries were associated with a broad overall measure of gender inequality, which reflected women's status across several domains including political, economic, legal, marriage and family, and discrimination.

Although women in the U.S. may not be subject to the same degree of harsh treatment and subordinate status that still plagues women in many parts of the world, the U.S. is far from achieving gender equity. In 2014, the median annual earnings for U.S. women was 78.6% of men's annual earning among full-time year-round workers (Hegewisch and DuMonthier 2016). In addition to this wage gap, women remain severely underrepresented in leadership positions in business and government. In 2016, U.S. women held only 4.4% of CEO positions at S&P 500 companies (Catalyst 2016), 19.4% of seats in the U.S. congress, and 24.6% of seats in state legislatures (Center for American Women in Politics 2016b). If the findings from comparative work in the developing world translate to the U.S. context, we would expect this persistent gender inequality in the U.S. to have implications for infant mortality. Levels of gender inequality vary substantially across U.S. states and may be associated with variation in infant mortality rates.

However, research on this topic in the U.S. has been quite limited. Only two studies examined the relationship between women's status and IMRs at the state level in the U.S. (Kawachi et al. 1999; Koenen et al. 2006). Both studies found lower IMRs in U.S. states where women had greater economic autonomy and reproductive rights. Neither study identified a relationship between the political dimension of women's status and IMRs net of controls for state racial composition and economic conditions.

The lack of a clearly established and robust relationship between political gender inequality and IMRs in these U.S. studies is particularly puzzling given the evidence of such a relationship in the developing world. This discrepancy may be attributable to differences in the conceptualization and measurement of political gender inequality. Studies in the developing world that identified a relationship between political gender inequality and IMRs focused specifically on women's legislative representation and measured the percentage of national parliamentary seats that were occupied by women (e.g. Quamruzzaman and Lange, 2016; Swiss et al., 2012). Both prior U.S. studies employed an index of "political participation" that combined women's representation in elected offices with levels of women's voter registration and women's voter turn-out. Because these women's voting measures were not relative to men's voting measures they are not necessarily reflective of gender inequality per se; political participation may simply be low among all state residents regardless of gender. The extent to which these voting measures correspond to meaningful differences in political power is also unclear. For these reasons the present study follows the more straightforward approach of research in the developing world and examines women's representation in U.S. state legislatures as the key indicator of state-level political gender inequality.

There are two principal theoretical rationales for why there should be a relationship between women's political representation and infant mortality. The first rationale relies on an ecosocial perspective to connect gender inequality in a given social

context to maternal and infant health. Nancy Krieger's (2014, 2001) ecosocial theory describes the processes by which people embody (i.e., literally biologically incorporate) social inequality. In the ecosocial model, social patterning of disease happens when exploitative/oppressive social relations structure exposure to living and working conditions, and to the exercise of civil, political, social, economic and cultural rights, which leads to the biological expression of discriminatory social systems. This embodiment occurs through multiple pathways including: economic and social deprivation, toxic/hazardous living conditions, socially inflicted trauma and inadequate healthcare (Krieger 2014, 2001). The ecosocial perspective undergirds a new line of research that uses the political representation of blacks as an indicator of structural racism (Lukachko, Hatzenbuehler, and Keyes 2014) and the political representation of women as an indicator of structural gender inequality (e.g. Backhans et al., 2007, 2009).

From an ecosocial perspective, underrepresentation of women in positions of political power is part of a system of gender inequality that may dramatically shape not only women's life chances, but also their health and the health of their children. As Koenen et al. (2006) point out, infants can literally embody the social status of their mothers through both the biological connection *in utero* and the intimate social connection postpartum. One specific example of a relevant ecosocial pathway that is supported by current empirical research is the connection between economic gender inequality, domestic violence, and infant mortality. The gender wage gap in local labor

markets is positively associated with violence against women (Aizer 2010) and exposure to domestic violence is a serious risk factor for infant mortality, low birth weight and preterm birth (Saifuddin, Koenig, and Stephenson 2006; Shah and Shah 2010).

Several complementary theories and findings from health disparities research suggest other potential pathways through which gender inequality may be embodied by women and their infants. For example, living in a context of greater gender inequality may harm women's health before and during pregnancy through psychosocial pathways related to low social status and relative deprivation (Subramanyam et al. 2009; Wilkinson 1997; Wilkinson and Pickett 2007). Similarly, increased stress exposure has been identified as an important mechanism linking race and gender to poor health (Sternthal, Slopen, and Williams 2011; Thoits 2010), and maternal stress is a known contributor to pregnancy complications and poor infant outcomes (Mulder et al. 2003). Finally, perceived discrimination has been linked to infant mortality and low birth weight in racial and ethnic minority populations (Collins et al. 2004; Krieger et al. 2013; Lauderdale 2006). Although these pathways have been tested mainly in relation to racial and socioeconomic inequality, they are nonetheless potentially important mechanisms for the nascent line of research on contextual gender inequality to explore.

The second theoretical rationale for the relationship between women's political representation and infant mortality is broadly consistent with the ecosocial perspective, but emphasizes political processes and more direct policy pathways. Many scholars

have argued that the higher women's status in society, the greater their influence over social policy, which results in more investments being made in education, healthcare, nutrition and other social programs that reduce infant mortality (Boehmer and Williamson 1996; Koenen et al. 2006; World Bank 2003). Miller's (2008) historical analysis supports this type of pathway by providing evidence that the extension of suffrage rights to U.S. women in 1920 prompted a large increase in public health spending which reduced child mortality. Prior research has also shown that women legislators are more likely than their male counterparts to make budgetary allocations and enact policies that promote healthcare, women's and children's issues, social services, education, welfare spending and anti-discrimination policy (Bolzendahl 2009; Bolzendahl and Brooks 2007; Bratton and Haynie 1999; Caiazza 2004; Little, Dunn, and Deen 2001; Taylor-Robinson and Heath 2003; Thomas 1991), all of which could potentially improve infant health outcomes. Indeed, research has found that policies that support Medicaid, maternal and infant care programs, subsidized family planning, and access to abortion are associated with reductions in infant mortality and improvements in maternal and child well-being (Bailey, Guldi, and Hershbein 2013; Grossman and Jacobowitz 1981).

Although testing specific social policy mechanisms is outside the scope of the present study, the crucial first step undertaken here is to reassess whether there is, in fact, a relationship between women's political representation and IMRs in the U.S. The

present study also extends prior research by using longitudinal data. Both previous studies of IMR and women's status in the U.S. (Kawachi et al. 1999; Koenen et al. 2006) employed cross-sectional designs, which do not allow for the observation of trends or changes across time. Cross-sectional designs also introduce the possibility of confounding by unobserved state-level characteristics that may pose competing explanations for the patterns observed. The longitudinal data used in the current study permit an analytic approach that models within-state changes in the key variables over time. This means that the influence of unobserved time-invariant state-level characteristics can be both modeled and excluded, thus more precisely pinpointing the relationship between women's political representation and state infant mortality rates. The present study uses U.S. state-level data measured at 12 time points across a period of 22 years to (1) investigate the relationship between the political representation of women in state legislatures and state infant mortality rates, and (2) project the U.S. population level costs, in terms of infant mortality, associated with gender inequality in state legislatures in 2012.

## **2.2. Methods**

### **2.2.1. Sample**

Data for this study are compiled from several administrative sources listed below. Each variable is measured for each of the 50 U.S. states every two years from

1990 to 2012. This yields a total sample of 600 state-years (50 states times 12 observation periods).

### **2.2.2. Measures**

*Infant Mortality Rate.* Infant mortality rates are the number of deaths occurring to infants under 1 year of age per 1,000 live births. These rates were recorded by the CDC/NCHS and compiled at the state level by the Kids Count Data Center (Annie E. Casey Foundation 2016). They are included in models as rates per 1,000.

*Percent Women in State Legislature.* State-level political gender inequality is measured as the percentage of state legislature seats that are occupied by women (i.e., (# women legislators / total # legislators)\*100). These percentages are provided by the Center for American Women In Politics (Center for American Women in Politics 2016a).

*Control Variables.* Multivariate models include a linear term for year to account for any secular trend in IMR across time. Controlling for the secular trend is crucial because IMRs have been declining in the U.S. across time for the past century (with the exception of a brief period between 2000-2005) and this decline has been attributed to a variety of factors not measured in the present analysis including: improvements in sanitation, medical technology and diffusion, and access to prenatal care (U.S. Department of Health and Human Services, 2014; see also Wang et al, 2014). Additional controls include the following time-varying state-level characteristics that may be associated with IMRs and women's political representation. The percentage of the state

population that is below the federal poverty line in each year is based on data from the U.S. Census Bureau (United States Census Bureau 2016a). The percentage of the state population with a high school degree or higher in each year is calculated using IPUMS-CPS data (Flood et al. 2015). The median household income for each state is from the U.S. Census Bureau's Small Area Income and Poverty Estimates (SAIPE) (United States Census Bureau 2016c). The timing of the SAIPE measures does not exactly coincide with the other measures for the first three of the twelve time points in this study. Therefore, the 1990 median household income measure uses the 1989 SAIPE estimate, the 1994 measure uses 1993 SAIPE estimate, and the 1992 measure uses the mean of the 1989 and 1993 SAIPE estimates. Median household income is included in models as thousands of 1999 U.S. dollars. State income inequality is measured in each year using the Gini coefficient provided by Frank (2013) (also see Frank, 2014 for details). State racial composition is measured by the percentage of the state population that is nonwhite in each year based on census population estimates (United States Census Bureau 2016b). Some random-effects models also control for regional differences by including a time-invariant dummy variable indicating whether each state is considered part of the South based on the U.S. Census Bureau's region definitions (United States Census Bureau 2015).

A supplemental model (Model 1 in Appendix A Table 13) also includes a series of dummy variables indicating partisan control of each state legislature as either

republican majority, democratic majority, or split, in each year (National Conference of State Legislatures 2016). This strategy parallels that of cross-national work which typically includes a dummy variable representing left-party in power (Bolzendahl 2009; Quamruzzaman and Lange 2016; Swiss et al. 2012). The partisan control variables are not included in the main models in Table 3 because they do not have a statistically significant relationship to IMRs and because they require the exclusion of all observations for Nebraska due to its non-partisan legislature. An additional supplemental model (Model 2 in Appendix A Table 13) adds a control for the political party affiliation of the sitting U.S. President (Rodriguez, Bound, and Geronimus 2013).

### **2.2.3. Analytic Methods**

To predict the change in state infant mortality rates given changes in the proportion of women in state legislatures, I estimate both fixed-effects models and random-effects models. Because there are many unobserved state-level characteristics (beyond those included as model controls) that could potentially be related to both women's political representation and IMR, a fixed-effects model is useful for removing the effects of these potential confounders. This is accomplished by using each state as its own control, and therefore only within-state change over time is modeled. However, if there is no state-level confounding then using fixed-effects can lead to an unnecessary loss of statistical power and unbiased estimates can be more efficiently estimated with a random-effects model, which models IMR using both within- and between-state

variation in women's political representation. The Hausman Specification Test ( $\chi^2(7)=10.30, P=0.172$ ) indicates there are no statistically significant differences between the coefficients estimated by the fixed- and random-effects models, suggesting that the random-effects model is appropriate (Allison 2009). Therefore, results are discussed in detail only for the random-effects models, but the fixed-effects parameter estimates are also included in Table 3 for comparison. All random-effects models shown use robust clustered standard errors to adjust for any cross-sectional heteroskedasticity and within-panel serial correlation of errors.

The models take the following forms:

#### Random-Effects Models

$$y_{it} = b_0 + \sum_{k=1}^K b_k x_{ikt} + u_i + e_{it},$$

where  $y_{it}$  is the IMR for state  $i$  in year  $t$ , and  $b_0$  is an intercept,  $x_{ikt}$  is the  $k^{\text{th}}$  covariate for state  $i$  at time  $t$  and  $b_k$  is its effect,  $u_i$  is a time-invariant random effect for state  $i$ , such that  $u \sim N(0, \tau)$ , and  $e_{it}$  is a state- and time-specific residual that follows the usual OLS assumptions of homoskedasticity and independence.

#### Fixed-Effects Models

$$y_{it}^* = \sum_{k=1}^K b_k x_{ikt}^* + e_{it},$$

where  $y_{it}$  is the IMR for state  $i$  in year  $t$ , with  $y_{it}^* = y_{it} - \bar{y}_i$  and  $b_0$  is an intercept,  $x_{ikt}$  is the  $k^{th}$  covariate for state  $i$  at time  $t$ , with  $x_{ikt}^* = x_{ikt} - \bar{x}_{ik}$ , and  $e_{it}$  is a state-and time-specific residual that follows the usual OLS assumptions.

A final analytical consideration is the timing of key measures. Specifically, should models include a time lag between the measures of political gender inequality and infant mortality outcomes? And if so, what is the appropriate length of the lag? The results shown are from models using contemporaneous values of the variables rather than a lag, because this strategy produces reliable, conservative estimates when the true lag is unknown (e.g. see Lynch, 2011). Furthermore, including a lag reduces the amount of usable data and introduces issues with interpretation and timing of other covariates. Nonetheless, as a sensitivity test, I estimate additional models measuring the proportion of women in the states legislatures with a two-year, a four-year, and a six-year lag (see Appendix A Table 14). The results are quite similar. For the two and four year lag periods, the estimated effects of women's political representation remain statistically significant and within a range of 4 thousandths of a point of results presented here.

To test for multicollinearity, I estimate a series of OLS models and calculate variance inflation factors (VIFs) from these models for each covariate in the full model (included in Appendix A Table 15). The VIF estimate for the key independent variable, legislature % women, is 1.70. All VIFs are below the most common threshold of 10 and

the more stringent threshold of 5, indicating that multicollinearity is not a cause for concern in the analysis.

### 2.3. Results

Table 1 shows descriptive statistics for the key variables in the analysis. There is substantial variation between states and within states across time in both the IMR and the proportion of legislature seats occupied by women. The average proportion of women in state legislatures in the U.S. increased modestly from 17% in 1990 to 23.9% in 2012. This is still far below parity with men. States with the lowest percentages of

**Table 1. U.S. State-Level Descriptive Statistics, 1990-2012**

1990				
	<u>Mean</u>	<u>SD</u>	<u>Min</u>	<u>Max</u>
Percent of state legislature seats occupied by women	17.0%	7.7%	2.1%	33.3%
IMR	9.06	1.42	6.20	12.40
Percent of population below poverty line	13.1%	4.2%	6.0%	25.7%
Percent of population nonwhite	19.0%	13.5%	2.0%	69.0%
Median household income (thousands of dollars)	38.0	6.2	26.7	53.2
Percent of population with high school degree or higher	78.5%	5.8%	65.9%	87.6%
Gini coefficient	0.565	0.025	0.531	0.671
2012				
Percent of state legislature seats occupied by women	23.9%	6.7%	10.0%	40.0%
IMR	6.16	1.25	4.20	8.90
Percent of population below poverty line	14.2%	3.4%	8.1%	22.0%
Percent of population nonwhite	26.5%	15.0%	5.0%	74.0%
Median household income (thousands of dollars)	38.3	6.3	27.5	52.7
Percent of population with high school degree or higher	89.1%	2.9%	81.8%	94.0%
Gini coefficient	0.617	0.037	0.548	0.711
Longitudinal 1990-2012				
Percent of state legislature seats occupied by women				
Overall	21.7%	7.6%	2.1%	40.8%
Between-state		6.7%	8.5%	35.8%
Within- state		3.5%	9.9%	32.6%

IMR				
Overall	7.24	1.57	3.70	12.40
Between-state		1.22	5.20	10.50
Within-state		1.01	4.23	11.03
Percent of population below poverty line				
Overall	12.7%	3.6%	4.5%	25.7%
Between-state		3.1%	6.8%	20.3%
Within-state		1.8%	7.3%	18.5%
Percent of population nonwhite				
Overall	22.40%	14.40%	2.00%	74%
Between-state		14.30%	3.10%	72.30%
Within- state		2.70%	11.50%	33.50%
Median household income (thousands of dollars)				
Overall	39.1	6.2	26.2	56.7
Between-state		5.8	29.0	51.9
Within-state		2.3	33.2	44.8
Percent of population with high school degree or higher				
Overall	84.8%	5.3%	65.9%	94.4%
Between-state		3.9%	77.6%	91.0%
Within-state		3.6%	72.3%	94.8%
Gini coefficient				
Overall	0.586	0.036	0.526	0.711
Between-state		0.024	0.550	0.643
Within-state		0.026	0.532	0.692

women in their legislatures in 2012 include: South Carolina (10%), Louisiana (11.1%), Oklahoma (12.8%), Alabama (13.6%), North Dakota (15.6%), and Wyoming (15.6%).

States with the highest percentages of women in their legislatures in 2012 include:

Colorado (40%), Vermont (38.9%), Hawaii (35.5%), Arizona (33.3%), and Minnesota

(32.8%). The average state IMR declined from 9.06 per 1,000 in 1990 to 6.16 per 1,000 in

2012. Although this is a substantial decline, the rate remains among the highest in the

developed world. States with the highest IMRs in 2012 include: Alabama (8.9),

Mississippi (8.8), South Dakota (8.7), and Louisiana (8.2). States with the lowest IMRs in

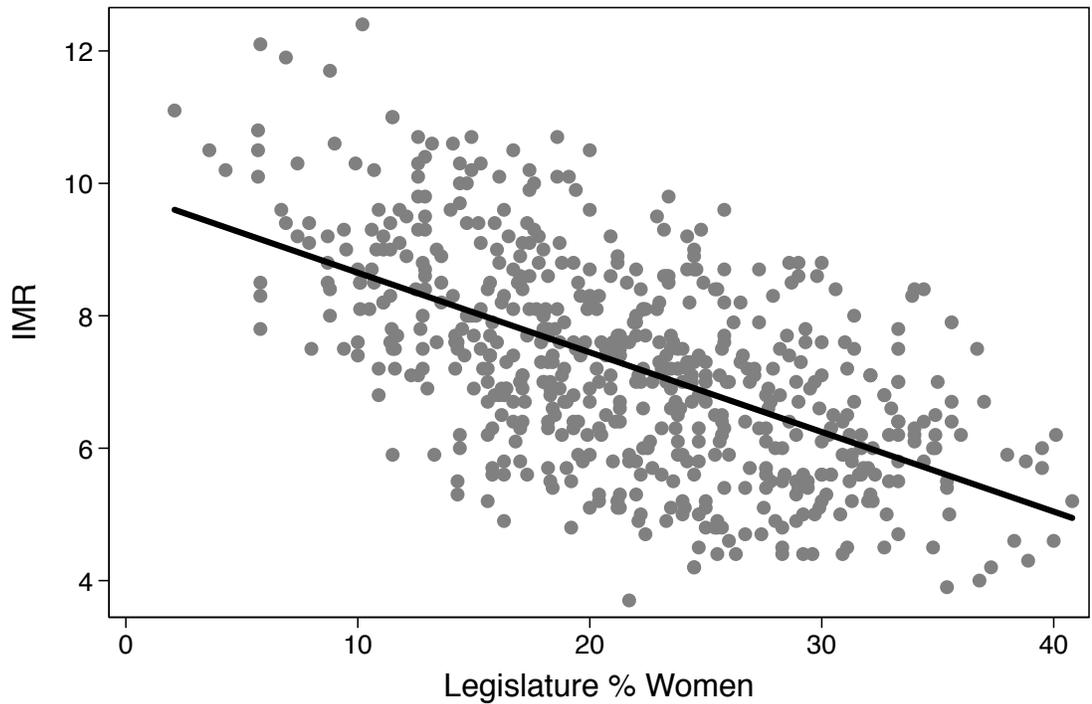
2012 include: New Hampshire (4.2), Massachusetts (4.2), Vermont (4.3), New Jersey (4.4), and California (4.5).

Table 2 shows the correlations between all variables used in the main analyses. State IMR has a strong negative correlation with: the percent of the state population with a high school degree or higher, the proportion of state legislators that are women, median household income, and year. IMR is positively correlated with the southern region dummy variable and with the poverty rate.

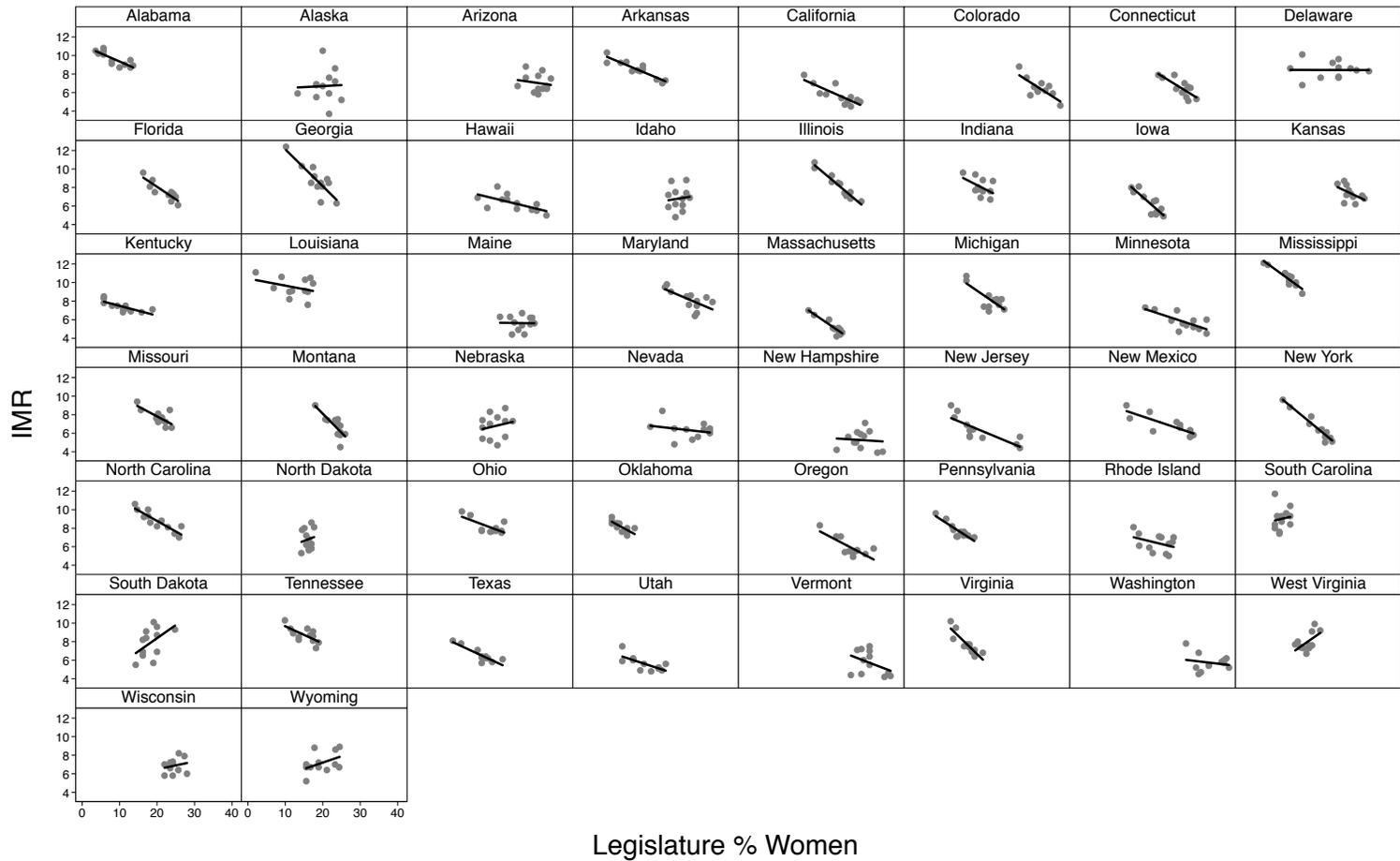
**Table 2. Zero-Order Correlations Between Key Variables**

	IMR	Legislature % Women	Poverty	Nonwhite	Median Income	% HS Degree	Gini
Legislature % Women	-0.578						
Poverty	0.397	-0.388					
Nonwhite	0.074	0.004	0.316				
Median Income	-0.444	0.456	-0.699	0.204			
% HS degree	-0.646	0.547	-0.549	-0.170	0.468		
Gini	-0.219	0.054	0.210	0.312	0.030	0.221	
Southern Region	0.529	-0.497	0.451	0.257	-0.366	-0.517	0.015
Year	-0.490	0.257	0.009	0.165	0.153	0.590	0.474

The bivariate relationship between the gender composition of state legislatures and IMRs is illustrated in Figure 1. The scatter plot, which includes values for all 50 states at each of the 12 time points, shows a clear negative linear relationship. A similar pattern is also evident within the majority of states over time (Figure 2). Thus, in a bivariate context, higher proportions of women in state legislatures are associated with lower infant mortality rates. (Scatter plots illustrating the bivariate relationships between IMRs and other key independent variables are included in Appendix A Figure A1.)



**Figure 1. IMR by Percentage Women in Legislature, U.S. States, 1990-2012 (N=600)**



**Figure 2. Within-State Relationships Between IMR and Women in Legislature, 1990-2012**

Table 3 shows the results from the longitudinal regression models predicting state infant mortality rates. Four different models are shown in both a random-effects framework and a fixed-effects framework (for a total of eight models). Model 1 includes only women's political representation. Model 2 adds a linear term for the year to account for any possible secular trend in IMR across time. Model 3 adds controls for the poverty rate and southern region. Model 4 is the fully adjusted model, which includes additional controls for the socioeconomic and racial characteristics of the states.

Focusing on the random-effects models, the findings are as follows.

Model R1 results show a strong negative relationship between women's political representation and IMRs. Model R2 results show a trend in decreasing IMRs across time between 1990 and 2012, such that each consecutive two-year period is associated with an average decrease of 0.099 infant deaths per 1,000 ( $p < .001$ ). The effect size of women's political representation is reduced somewhat when the time trend is introduced in model R2, but the effect remains substantial and statistically significant. Model R3 also shows that women's political representation is negatively related to infant mortality. For every one-percentage point increase in state legislature seats occupied by women, infant mortality rates decrease by 0.034 per 1,000 ( $p < .01$ ). This result implies a sizable predicted difference of 1.32 infant deaths per 1,000 live births between the states with the lowest levels of women political representation in the sample and those with the highest. Model R3 results also show that higher state poverty levels and being located in the

South are associated with higher levels of infant mortality. For every one-percentage point increase in the population living below the poverty line, IMR increases by 0.032 per 1,000 ( $p < .05$ ). This result implies a predicted difference of approximately 0.68 infant deaths per 1,000 live births between the poorest and richest states in the sample.

Southern states have predicted IMRs higher than non-southern states by 1.401 per 1,000 ( $p < .001$ ).

Finally, model R4 results show that adding additional controls for state socioeconomic and racial characteristics has a negligible impact on the estimated effect of women's political representation. The coefficient for legislature % women, remains statistically significant and quite close in size. Educational attainment, median household income, Gini coefficient, and % nonwhite are not statistically significant predictors of state infant mortality rates net of the other controls already in the model. When the additional controls are added the effect of the poverty rate also becomes non-significant. The fully adjusted model helps to rule out potential mediators and/or confounders, but adding these additional control variables did not improve model fit, as evidenced by both the Wald test for the random-effects models and the F-test for the fixed-effects models (Wald  $\chi^2(4) = 1.50$ ,  $P = 0.872$ ;  $F(4, 534) = 1.32$ ,  $P = 0.261$ ). Therefore, the best fitting and most parsimonious model (R3) is used for the population estimates that follow, but estimates produced using model R4 do not differ meaningfully from those presented here. Considering all measured covariates and across all model specifications,

higher percentages of women in state legislatures are associated with reduced IMRs.

The results from the fixed-effects models (F1-F4) are quite similar to random-effects results, suggesting that unobserved state-level fixed characteristics cannot account for the relationship between women's representation and IMRs.

**Table 3. Results of Longitudinal Regression Models Predicting U.S. State Infant Mortality Rates, 1990-2012**

	Random-Effects Models				Fixed-Effects Models			
	Model R1	Model R2	Model R3	Model R4	Model F1	Model F2	Model F3	Model F4
Legislature %								
Women	-0.135*** (0.01)	-0.045*** (0.01)	-0.034** (0.01)	-0.032* (0.01)	-0.141*** (0.01)	-0.031** (0.01)	-0.029** (0.01)	-0.025** (0.01)
Year		-0.099*** (0.01)	-0.102*** (0.01)	-0.096*** (0.01)		-0.103*** (0.00)	-0.104*** (0.00)	-0.089*** (0.01)
% Poverty			0.032* (0.02)	0.028 (0.03)			0.025 (0.02)	0.034 (0.02)
Southern Region (ref: non-south)			1.401*** (0.28)	1.351*** (0.28)				
% HS Degree or Higher				-0.019 (0.02)				-0.026 (0.02)
Median Income				0.002 (0.02)				0.017 (0.02)
Gini				1.314 (2.29)				1.876 (1.37)
% nonwhite				-0.003 (0.01)				-0.031 (0.02)
Constant	10.175*** (0.34)	206.210*** (13.85)	211.922*** (13.83)	200.937*** (24.91)	10.298*** (0.23)	214.188*** (9.44)	215.136*** (9.45)	187.523*** (25.61)
n	600	600	600	600	600	600	600	600
Model R2: Overall	0.334	0.403	0.579	0.576				
Between-state	0.402	0.402	0.571	0.563				
Within-state	0.244	0.590	0.593	0.595	0.244	0.591	0.593	0.597

Notes: \* p<.05, \*\* p<.01, \*\*\*p<.001, two-tailed tests. Standard errors are included in parentheses.

Given the strong negative relationship between the political representation of women and infant mortality, how can we quantify the costs to the U.S. population of the persistent underrepresentation of women in state governments? To address this question, I use random-effects model R3 to predict and compare: (1) the IMR in 2012 for a state at the 2012 observed mean level of women's political representation (23.886% of legislature seats occupied by women) and (2) the IMR a state would have if women and men were at parity (i.e. 50% of legislature seats occupied by women). At the 2012 mean level of women's political representation, the model predicts an IMR of 6.04 per 1,000 (SE= 0.120; 88% CI =5.85, 6.23). At parity, the model predicted IMR is 5.16 per 1,000 (SE= 0.347; 88% CI= 4.62, 5.70). I construct an 88% confidence interval to compare these point estimates because, based on the ratio of the standard errors, a z-score of 1.54 (corresponding to roughly an 88% CI) is appropriate for determining whether the difference between these estimates is statistically significant at the .05 level. A traditional 95% CI would be too conservative a test and would likely result in a type I error (Schenker and Gentleman 2001). Furthermore, an independent samples z-test of the difference indicates that the difference of 0.88 per 1,000 between the predicted IMR point estimates is statistically significant at the .05 level ( $z=31.54$ ,  $p<0.0001$ ).

Beyond establishing statistical significance, some additional calculations are warranted to demonstrate the substantive importance of this difference. The total number of registered births in the US in 2012 was 3,952,841 and the number of infant

deaths in 2012 recorded by the CDC was 23,629 (Annie E. Casey Foundation, 2016).

Applying the predicted IMRs from the regressions to the number of registered births yields the following estimates:

23,875 infant deaths at observed mean (88% CI=23,124, 24,626)

20,397 infant deaths at parity (88% CI=18,262, 22,531)

Subtracting these estimates suggests 3,478 excess infant deaths in the U.S. attributable to gender inequality in political representation in 2012. This estimate implies a hypothetical 14.6% reduction in infant mortality in 2012 if the political representation of women in state legislatures had been equal to that of men.

## **2.4. Discussion**

This study is the first to identify a strong negative relationship between the political representation of women in state governments and state infant mortality rates in the United States. This relationship is evident both within individual states as the representation of women changes across time, and between states with high vs. low levels of women's representation. The fact that consistent findings were observed in both the random-effect and fixed-effects models indicates that the relationship between women's political representation and IMRs cannot be attributed solely to differences in unobserved time-invariant characteristics of U.S. states. The relationship also holds with a number of time-varying covariates statistically controlled. The results suggest that political gender inequality is of equal or greater importance as a determinant of state

infant mortality rates than state poverty levels. This fact highlights the substantive importance of these findings given the longstanding recognition of the strong link between poverty and infant mortality (Center for Disease Control and Prevention 1995; Gortmaker 1979).

One potential competing explanation for the patterns observed here might be that the relationship between women's representation and infant mortality is due primarily to partisan differences rather than gender representation. This explanation would be plausible if, for example, female legislators were overwhelmingly affiliated with the Democratic Party and if Democratic control of state legislatures was linked to lower infant mortality rates. However, the supplemental analyses conducted indicate that majority political party in the legislature is not a statistically significant predictor of state infant mortality rates (see Appendix A Table 13). This result is consistent with studies in the developing world that find women's representation is not simply a proxy for left party power (Bolzendahl 2009; Quamruzzaman and Lange 2016; Swiss et al. 2012). Thus, partisan composition of legislatures cannot account for the relationship between women's political representation and infant mortality.

Another partisan consideration is the political party affiliation of the U.S. President. A recent study by Rodriguez and colleagues (2013) found that IMRs were higher during Republican administrations than Democratic administrations. Therefore, a supplemental analysis was conducted with additional controls for the President's

party affiliation (see Appendix A Table 13). The results show that the main findings of this study are robust: the relationship between women's political representation and IMR remains negative and statistically significant.

The nonsignificant coefficients for educational attainment in the fully adjusted models (R4 and F4 in Table 3) may initially seem odd given the widespread recognition of education as an important determinant of health in general (Braveman et al. 2010; Ross and Wu 1995) and infant mortality specifically (Shandra et al. 2004; Singh and Yu 1995; Wang et al. 2014). However, this lack of significance occurs primarily because educational attainment levels increased each year in every U.S. state during the study period, and these increases are captured by the linear term for year which is already entered into the models as a control for the secular trend in IMR (see Appendix A Table 16). In random-effects models there remains some significant effect of educational attainment on IMRs after controlling for year (due to the inclusion of between-state variation in the random-effects model), but this effect is roughly one quarter its original size and it loses statistical significance when additional controls are introduced (see Appendix A Table 17). These findings are consistent with Wang et al.'s (2014) conclusion that secular trends account for a much larger portion of the global decline in child mortality than increases in educational attainment do.

The results of the present study do not necessarily imply that educational attainment is unimportant for infant mortality in the U.S.. Rather, increasing levels of

educational attainment are likely to be one of many factors (including technological/medical advancements and diffusion) that contribute to the declining secular trend in IMRs observed in the U.S.. Furthermore, the present study is conducted at the state level and therefore it does not address the relationship between education and infant deaths at the individual level. A non-significant effect of state-level educational attainment does not contradict the finding that individual women with higher levels of education have more favorable birth outcomes than those with lower levels (e.g. Singh and Yu, 1995).

The main limitation of the present study is that—like the prior studies of gender inequality and IMRs in the United States (Kawachi et al. 1999; Koenen et al. 2006)—the analysis is conducted at the ecological level and, therefore, individual level inferences cannot be made. Future multilevel analyses are needed to estimate the risks to individual women and children posed by living in a social context characterized by gender inequality. Multilevel analyses will also allow for investigation into potential mediators and mechanisms.

This study does not attempt to identify the mechanisms by which greater political representation of women may reduce infant mortality. One plausible mechanism is that women who are elected may use their political power in ways that are beneficial for women, children, the poor, and overall public health (Bolzendahl 2009; Bolzendahl and Brooks 2007; Bratton and Haynie 1999; Caiazza 2004; Little et al. 2001;

Taylor-Robinson and Heath 2003; Thomas 1991). The extent to which these policy differences explain the relationship between political representation and infant mortality remains unclear. Quamruzzaman and Lange (2016) found that health spending partially mediates the relationship between women's political representation and IMRs in developing countries, but it only accounts for a small portion of the relationship. Researchers should explore state-specific policies and social spending pattern to see if they can help explain this relationship. Investigating potential mechanisms is a major priority for future research in order to better understand why and how women's political representation influences infant mortality in the U.S.

In addition to exploring specific policy mechanisms, future research should focus on understanding upstream macro-social factors that comprise the gender structure of a given social context. The underrepresentation of women in state legislatures most likely reflects a broader climate of gender inequality that extends beyond the political realm to social, cultural, and economic conditions, all of which may also play a role in shaping women's health (Moss 2002). More theoretical and empirical work is needed to situate gender inequality in political representation with respect to indicators of other domains of structural gender inequality and to uncover the implications of such inequality for individual and population health.

Finally, it is important to note that as of 2016, no U.S. state legislature has ever had levels of women's representation equal to or exceeding that of men's. Therefore, the

impact of political equality in U.S. state legislatures remains hypothetical. The maximum level of women's representation observed in the sample is 40.8%, which means that predictions for conditions of equality (i.e. parity) are outside the range of the data and, therefore, relatively uncertain. Nevertheless, this study finds an IMR difference of 1.32 ( $p < 0.001$ ) per 1,000 across the range of inequality levels observed.

In conclusion, this study demonstrates that the importance of women's political empowerment to population health is not exclusive to the developing world. Gender inequality in U.S. state legislatures is associated with a substantial cost to the population in terms of infant mortality. According to the World Economic Forum, political empowerment is the domain in which the U.S. exhibits its highest levels of gender inequality. Although the U.S. ranked 28<sup>th</sup> out of 145 countries in overall gender equality in 2015, it ranked 72<sup>nd</sup> in women's political empowerment and 81<sup>st</sup> for the number of women in parliament (Hausmann, Tyson, and Zahidi 2015). Any national strategy directed at reducing infant mortality will be incomplete if it fails to consider political representation of women and gender inequality more broadly as key social determinants. Efforts to rectify the underrepresentation of women in state governments are essential both for social justice and for population health.

### **3. This Gender System Makes Me Sick! Structural Sexism and Health at Midlife**

*“Gender is an institutionalized system of social practices for constituting people as two significantly different categories, men and women, and organizing social relations of inequality on the basis of that difference.”*

(Ridgeway and Correll 2004:510)

Social inequality in the U.S. is sickening. Literally. Social hierarchies defined by race, class, gender and other axes of inequality powerfully influence individuals’ health and longevity (Berkman, Kawachi, and Glymour 2014). Much of our current understanding of how social systems of inequality shape health is based on three key types of research: (1) studies that examine how directly perceived experiences of discrimination, mistreatment or low status influence health; (2) studies of physician bias or discrimination within medical institutions; and (3) studies of the patterns/disparities in health outcomes across social categories representing relatively advantaged versus disadvantaged statuses.

Specifically in the case of gender inequality and health, this line of inquiry takes the form of studies on sexual harassment or perceived gender discrimination; investigations of how physicians and medical institutions fail to offer equitable, unbiased, appropriate medical care for women; and examinations of gender differences in health outcomes. Each of these approaches is vital to our understanding of gender inequality and health, but all have important limitations. The perceived discrimination

work captures only a small portion of the effects of gender inequality on health because so much of the persistent gender inequality in American society is either unperceived, and/or not conceptualized as unfair, unequal or discriminatory (Berk 1985; Ridgeway 2011; West and Zimmerman 1987). The literature on gender bias among physicians and medical institutions is able to describe the ways gender systems influence medical care and is not limited by individual perception. However, *health care* is not the same as health, and many people only interact with medical institutions after they become ill. The social conditions in which people live and work have a greater impact on population health than medical care (House 2015). Therefore, many of the important ways gender inequality shapes health lay outside the purview of this work.

Research in the gender differences in health tradition (perhaps the most robust area of gender and health scholarship) categorizes individuals as men or women<sup>2</sup> and then compares health outcomes across these gender categories. The resulting health gaps are often assumed to arise in large part from structural inequalities between men and women (i.e. features of the gender system), but the nature of these inequalities is often unclear and the extent to which biology plays a role in the observed health difference is a subject of ongoing interest and debate (Bird and Rieker 2008).

Furthermore, this approach treats gender as a static trait of individuals and generally

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<sup>2</sup> This gender categorization is typically indistinguishable from the sex categorization as male/female and is self-reported by respondents in the majority of large health surveys.

does not examine how inequality in gender systems varies across social contexts in ways that may influence the health of both men and women, although recent scholarship has called for increased efforts to contextualize the gender differences in health research (Read and Gorman 2010).

Thus, in the dominant traditions of gender, inequality, and health research the broader structural inequalities that characterize gender systems remain largely unmeasured. Therefore, I advance a *structural sexism and health* perspective that shifts the focus from comparisons between men and women on health outcomes to within-gender comparisons across levels of exposure to discriminatory gender systems. Toward that end, the primary contributions of the present study are to develop measures of structural sexism-- *defined as systematic gender inequality in power and resources*-- at the macro-, meso-, and micro-levels of the gender system in the U.S., and examine their relationships to the health of men and women in middle age.

### **3.1. Gender Discrimination and Health**

Although far more scholarship has been devoted to racial discrimination than to gender discrimination, a growing body of research has investigated the relationship between gender discrimination and health. This research has focused almost exclusively on how women's health is harmed by perceived discrimination or sexual harassment in the workplace and by gender bias in medical institutions. Self-reported experiences of perceived gender discrimination and harassment in the workplace have been linked to a

variety of negative mental and physical health outcomes among women including emotional distress, anxiety, depression, headache, gastrointestinal symptoms, and functional limitations (McDonald 2012; Pavalko, Mossakowski, and Hamilton 2003; Swanson 1999). This perceived discrimination and harassment work provides a valuable but incomplete picture of the effects of systemic gender inequality on health because the processes that create and reproduce unequal gender systems are often unperceived, or are not conceptualized as unfair or discriminatory (Berk 1985; Ridgeway 2011; West and Zimmerman 1987). Particularly in the workplace, gendered organizations and discriminatory practices shape the allocation of women and men to various positions and compensation levels in ways that are typically outside the awareness of individuals whose lives they affect (Acker 1990; Correll, Benard, and Paik 2007; Rivera 2017; Rivera and Tilcsik 2016). But regardless of individual awareness, the degree to which a society inequitably distributes valued resources and opportunities, or otherwise unequally treats individuals along gender lines, may have a powerful influence on health (Krieger 2014).

Research on gender discrimination or bias among physicians and medical institutions has shown that women patients are less likely than men patients to receive the most effective, advanced treatments and diagnostic procedures for a variety of health conditions (Arber et al. 2006; Chapman, Tashkin, and Pye 2001; McMurray RJ et al. 1991; Raine 2000). Studies have also found evidence of anti-woman gender bias in

medical education and textbooks (Andrikopoulou et al. 2013; K Alexanderson 1998) and Medicaid reimbursement rates roughly 30% lower for female-specific (versus male-specific) surgical procedures (Benoit, Ma, and Upperman 2017). Because this line of research has used multiple methods including content analysis and audit studies, it has been able to illuminate ways gender discrimination shapes health without being limited by individual perceptions. However, it is limited in scope, as *health care* is only one of many factors that contribute to health. Scholars and policy-makers alike are increasingly recognizing the primary importance of social factors as determinants of health (cf. Braveman, Egerter, and Williams 2011; U.S. Department of Health and Human Services 2010). The social conditions in which people live and work, and the social policies that shape them, have far greater impact on population health than medical care or health policy (Bradley et al. 2016; House 2015). Thus, although the labor and healthcare markets are key venues of gender discrimination, existing research in these domains cannot account for the myriad ways individual health may be affected through other social institutions and processes that constitute the gendered stratification system of a society.

### **3.2. Gender Differences (Gaps) in Health Outcomes**

The most longstanding and influential tradition of gender and health research is one that examines gender differences in health and mortality. This line of research, which began in earnest in the 1980's, compares rates of death and various illnesses

between men and women. The patterns of gender difference in health and mortality are now well documented (for thorough reviews see Bird and Rieker 2008; Read and Gorman 2010, 2011). In general, women live longer than men, but are more likely to suffer chronic illness and disability. This work measures gender (or, to be more precise, sex category) as an attribute of individuals and asks the question: *How and why do men and women differ in terms of health and mortality?* The observed differences between categories are understood to represent a combination of social, biological, and behavioral factors, and a central concern of this work is to disentangle the various explanations for patterns of difference.

A more structural perspective that begins with the concept of gender as a *social system* of difference and inequality, might instead ask the related, but distinct, question: *How does the inequitable gendered distribution of power, status, and resources that characterizes the gender structure of a society shape the health of its members?* While the existing research on gender differences in health is invaluable for producing knowledge of who is and is not healthy in American society, it cannot answer this question because exposure to discriminatory gender structures has remained unmeasured. Gaps between men and women on health outcomes can provide clues about how a gender system works, but do not provide all the information needed to understand how *structural sexism*, i.e. *the systematic gender inequality in power and resources manifest in a given gender system*, shapes health.

To illustrate this point, Figure 3 shows three different hypothetical scenarios, depicting possible relationships between structural sexism and health problems among men and among women. In each of these scenarios, the gap between men's and women's health problems at the mean level of sexism exposure is identical. The first panel shows a scenario in which structural sexism has no effect on men's or women's health. In this case, the observed gender gap in health problems could be due to biology or some other unobserved factor. The second panel illustrates a scenario with the same average gender gap in health problems as the first panel, but structural sexism is related to health problems in opposite directions among women and men. In this scenario, gender relations are best thought of as zero-sum— higher levels of sexism result in increasing benefit to the dominant group (men) and increasing harm to the subordinate group (women). This type of pattern is one we might expect based on conflict theory and classical gender/sex stratification perspectives (Blumberg 1984; Chafetz 1984; Collins 1975, 1971). The third panel shows a scenario in which the gender gap is again the same, but both men's and women's health are harmed by higher levels of structural sexism. Such a scenario is consistent with modern feminist studies of masculinities and men's health which suggest that patriarchal social systems foster a toxic culture that harms men as well as women (Connell 2005, 2012; Courtenay 2000). Similarly, some recent theories of structural inequalities and health suggest a pattern of universal harm because inequality undermines the social fabric and makes the entire society less productive and

healthy (Lucas 2013; Wilkinson and Pickett 2011). Any of these three scenarios (and several others) is possible with the same observed gender gaps in health outcomes.

Thus, gender gaps alone are insufficient to determine how structural sexism influences health because they do not measure exposure; both men and women living in a certain social context are exposed to some level of gendered inequality, but its effects on their health may differ. Therefore, I advance a *structural sexism and health* perspective that complements and extends the existing gender and health literature by attempting to measure attributes of gender systems and focusing on within-gender comparisons across levels of exposure to discriminatory gender systems, rather than on gender gaps in health outcomes. This approach allows me to answer two research questions: (1) Does structural sexism have a statically significant effect on health among women and men? (2) If so, are the patterns in effects more consistent with a theory of zero-sum gender conflict or universal harm?

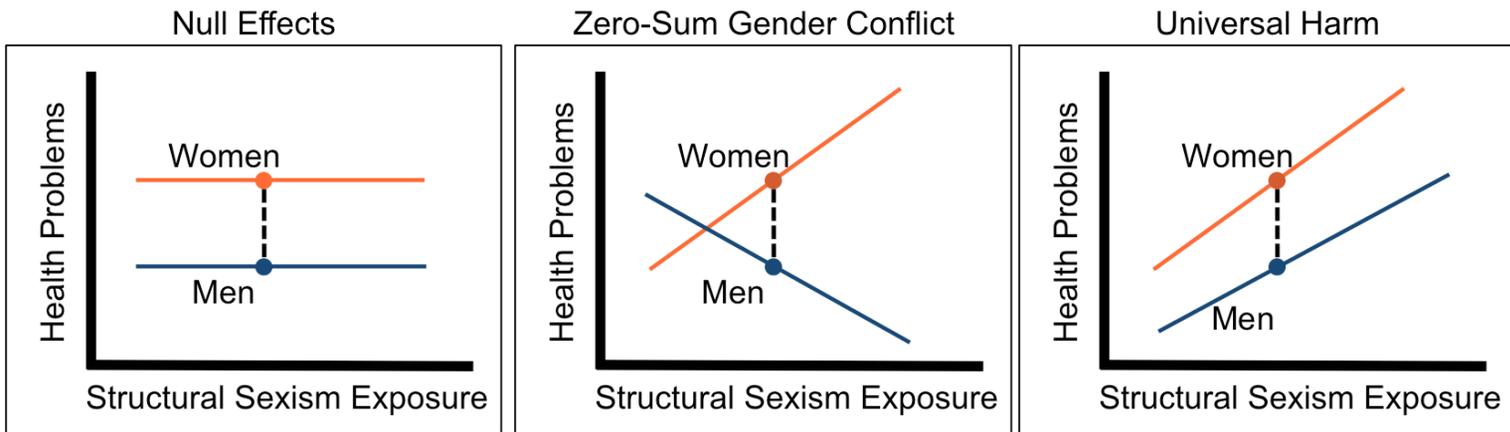


Figure 3. Hypothetical Relationships Between Structural Sexism and Health Among Women and Men

### ***3.3. Toward a Structural Sexism and Health Approach***

In developing a structural sexism approach, I draw on both the emerging structural racism and health literature and contemporary theories of gender as a multi-level social structure or system. Race and health research—similar to the gender and health research discussed above—has largely consisted of studies of health disparities across racial categories, physician bias and discrimination in medical institutions, and perceived discrimination. An immense amount of research has been devoted to these topics and the knowledge produced has been enormously influential. In particular, the Everyday Discrimination Scale (Williams et al. 1997) has been used in hundreds of studies to show that perceived discrimination is associated with a wide variety of physical and mental health problems (Pascoe and Richman 2009; Williams and Mohammed 2008).

However, a new structural racism and health literature has begun to grow out of a recognition that measures of perceived racial discrimination — while illuminating an important piece of the puzzle — stop short of capturing the full effects of racism on health because racism is not exclusively an interpersonal level phenomenon and because a large portion of racial discrimination goes unperceived (Bonilla-Silva 1997; Gee and Ford 2011). This new literature conceptualizes structural discrimination as a process that operates at the societal level to constrain the resources, opportunities, and well-being of disadvantaged groups (Hatzenbuehler et al. 2010; Link and Phelan 2001; Lukachko,

Hatzenbuehler, and Keyes 2014). By looking beyond individual actors and behaviors, this perspective highlights the discriminatory character of institutional arrangements. For instance, a recent study found that structural racism — as measured by state-level racial disparities in political representation, economic conditions, and juridical treatment — is associated with increased risk of myocardial infarction among Blacks in the United States and decreased risk among whites (Lukachko et al. 2014). Similar studies have linked other measures of state-level structural discrimination to increases in psychiatric disorders in lesbian, gay and bi-sexual populations (Hatzenbuehler et al. 2010), to higher rates of infant mortality among African Americans (Chae et al. 2018; Wallace et al. 2017), and to mortality rates among blacks, whites, males, and females (Lucas 2013). The application of this approach to the study of gender and health has been quite limited (Krieger 2014).

Contemporary gender scholarship conceptualizes gender not as an individual attribute, but as a multilevel structure (or system) of difference and inequality (Lorber 1994; Ridgeway and Correll 2004; Risman 2004). The gender system is relational and embodied (Connell 2012), and is expressed through institutions, interactions, and individuals in social processes that structure opportunities and constraints based on sex category (Risman 2004). Therefore, rather than focusing only on macro-level institutions as the structural racism and health literature has done, I propose a multilevel framework for structural sexism and health. A multilevel approach is particularly important given

how gender research and theory has highlighted the central role of interactional processes in the reproduction of unequal gender systems (Ridgeway 2011; Ridgeway and Smith-Lovin 1999; West and Zimmerman 1987). I use the term *structural sexism* to refer to *the systematic gender inequality in power and resources manifest in a given gender system*, and I argue that it can be observed at each level of the gender system. Figure 4 shows a conceptual model of structural sexism. Although structural sexism almost certainly enables more overt and intentional forms of gender based mistreatment such as interpersonal perceived gender discrimination and sexual harassment, structural sexism is considered conceptually distinct (Bonilla-Silva 1997; Krieger 2014; Lukachko et al. 2014).

Although most societies in history have been patriarchal, there is enormous variation in the extent of gender inequality in the distribution of power and resources across social contexts (Chafetz 1984). In this study, I seek to quantify the amount of structural sexism in various social contexts and leverage the observed variation across these contexts to understand the impact of structural sexism on the health of both men and women in the U.S. In the macro-, meso- and micro- subsections that follow, I describe in greater detail (1) how I conceptualize structural sexism at each level, (2) the domains in which I measure it for this particular study, and (3) the relevant prior research on how gender is related to health in each of these domains.

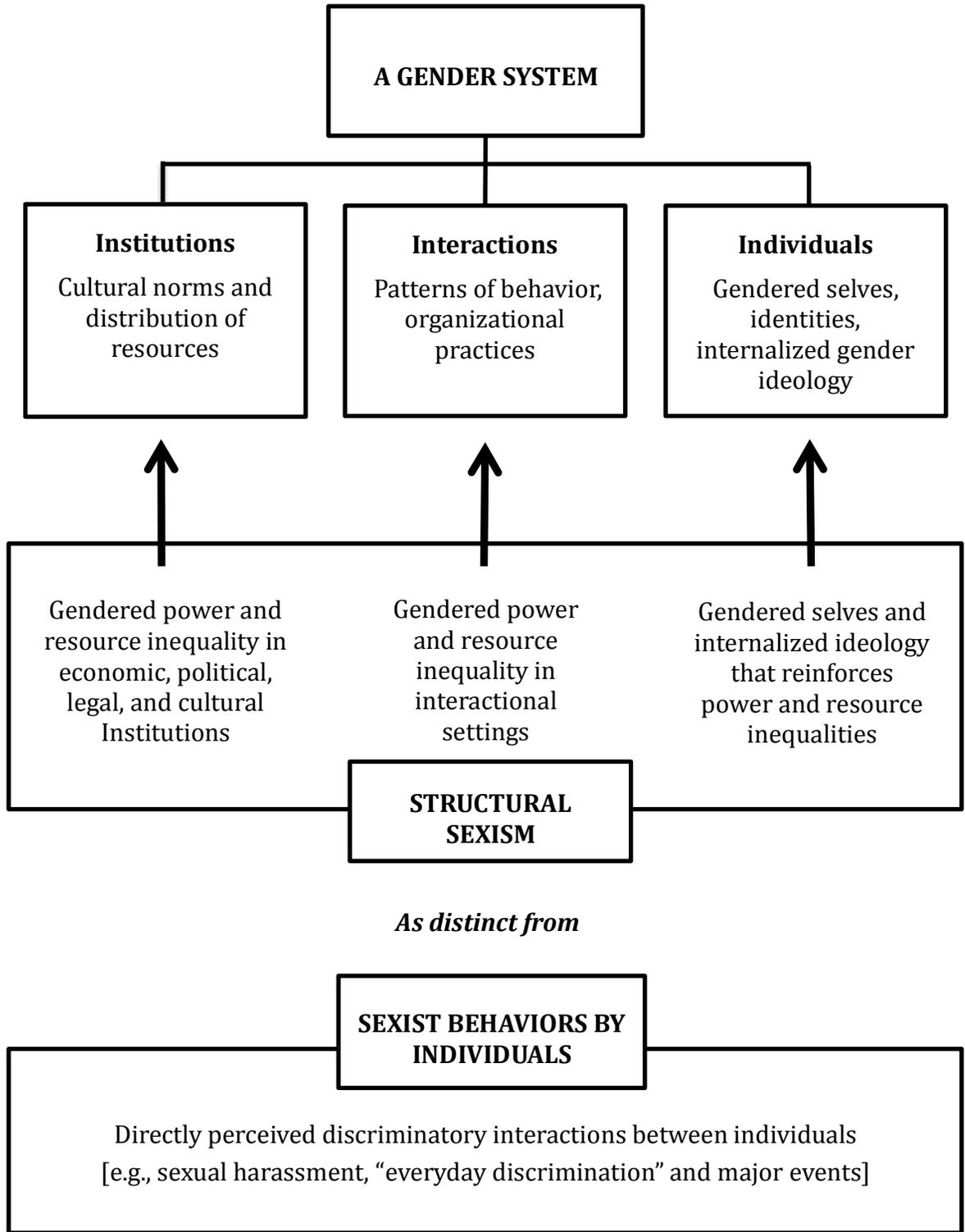


Figure 4. Conceptual Model of Structural Sexism

### **3.3.1. Macro-Structural Sexism**

At the macro-level, the gender system involves widespread cultural norms and the distribution of resources along gender lines in a society's major institutions (Ridgeway and Correll 2004; Risman 2004). Thus, macro-structural sexism refers to systematic gender inequality in power and resources favoring men within political, economic, and cultural institutions. While this can be thought of on both global and national scales, the present study is focused on structural sexism within the United States, and I will therefore examine U.S. state-level political, economic and cultural institutions. A recent study found that U.S. state-level characteristics were related to women's mortality rates, but did not include state-level measures of gender inequality (Montez, Zajacova, and Hayward 2016). The nascent structural racism and health literature on which I build has taken as the unit of study U.S. states and has demonstrated substantial variation in state-level discriminatory environments (Lukachko et al. 2014).

Very few studies have examined how state-level measures of gender inequality in the U.S. relate to population health. Three studies have used composite measures of "women's status," and they found low status is related to elevated state infant mortality rates (Kawachi et al. 1999; Koenen, Lincoln, and Appleton 2006), state mortality rates among women and men (Kawachi et al. 1999), and individual women's depressive symptoms (Chen et al. 2005). A fourth study found state-level measures of gender

inequality were related to individual men's mortality risk (Kavanagh, Shelley, and Stevenson 2017). Existing research has not examined the effects of U.S. state-level gender inequality (1) on chronic conditions and other individual physical health outcomes of both men and women, or (2) in conjunction with gender inequality at the micro and meso levels of a gender system.

### **3.3.2. Meso-Structural Sexism**

At the meso-level, the gender system involves interactions, patterns of behavior, and organizational practices (Ridgeway and Correll 2004; Risman 2004). Therefore, meso-structural sexism refers to the inequality in power and resources between men and women in interpersonal interaction. Gendered interactions occur regularly in the family, the workplace, and a variety of other social-relational settings, which are shaped by hegemonic cultural norms about gender (Connell 1987; Ridgeway and Smith-Lovin 1999). Men and women "do gender" in interaction when they orient their behavior toward these norms, thereby rendering social arrangements based on sex category normal and natural -- and therefore legitimate -- ways of organizing social life (West and Zimmerman 1987:146). Extensive research has described how the ways we "do gender" in the workplace and family perpetuate gender inequality in the division of labor (e.g. Berk 1985; Brines 1994; Coltrane 1989). Furthermore, the performance of gender-typed tasks and cultural expectations requiring women to be communal and team-oriented can result in women doing more undesirable or unrewarded work, thereby reducing their

likelihood of promotion (Babcock et al. 2017; Winslow 2010). This may not be perceived but may nonetheless shape women's health through direct and indirect pathways including stress and diminished access to material resources.

In addition to "doing gender," status processes have also been shown to perpetuate gender inequality in interaction in ways that individuals often do not perceive and even without individual discriminatory intent (Correll et al. 2007; Ridgeway 2011; Ridgeway and Smith-Lovin 1999). Ridgeway and Smith-Lovin (1999) describe the process as follows: men and women typically occupy structurally unequal positions when they interact; these status differences produce performance and perception differences that are confounded with gender difference; this reinforces gendered competence beliefs, thereby reproducing the gender system in interaction. Because men and women interact much more closely and frequently than people of different racial and class categories, measuring inequality at the interactional level has a unique importance for structural sexism compared to other forms of structural discrimination. Although there are a variety of contexts in which meso-structural sexism can be measured, I choose to focus on marriage for this study because of its primacy as site of interaction between men and women and its centrality to the reproduction of gender inequality (Orkin 1989).

There is a long history of studying the gendered health benefits of marriage going all the way back to Durkheim (1897). Most of this work shows that being married

is beneficial for health, but that men appear to benefit more than women (Berkman and Breslow 1983; Durkheim 1897; House, Landis, and Umberson 1988; Ross, Mirowsky, and Goldstein 1990; Umberson 1992; Umberson and Kroeger 2016), although some recent research calls into question the gender differences in the effects after accounting for marital quality (Carr and Springer 2010; Williams 2003). There is also a robust tradition of looking at inequality within marriage -- research that has focused mainly on the division of labor and the allocation of household chores (Berk 1985; Bianchi et al. 2000; Brines 1994; Coltrane 1989; Friedberg and Webb 2006). Surprisingly, very little research has connected this gender inequality in status and power within marriage directly to health. The few studies that have been done on this topic find that power imbalances within marriage are associated with higher levels of depression and psychiatric symptoms for the spouse that is in the low-power position (Bagarozzi 1990; Halloran 1998; Mirowsky 1985), and that the distribution of power in the average marriage favors the husband's mental health rather than the wife's (Mirowsky 1985). In the present analysis, I investigate whether interspousal inequality in power and resources -- one specific type of meso-structural sexism -- is related to women's and men's physical health.

### **3.3.3. Micro-Structural Sexism**

At the micro-level, the gender system involves gendered selves, identities, and internalized gender ideology (Ridgeway and Correll 2004; Risman 2004). Therefore, micro-structural sexism refers to gendered constructions of self and internalized gender ideology that undergird and reinforce gendered resource and power inequalities. This type of sexism is created through the processes of socialization, internalization, and identity work, and is embodied by individuals (Risman 2004). Although this is expressed in individuals, it can be thought of as structural because of its role in reproducing discriminatory gender structures. In this study, I will focus on individuals' gender role attitudes as reflections of internalized gender norms.

A large and growing literature has linked masculinity norms and gender ideology to men's health. In particular, studies show that traditional gender role beliefs as well as conformity to stereotypically masculine ideals, are linked to risk-taking, unhealthy behaviors (such as excessive substance use and violence) and healthcare avoidance (Courtenay 2000; Mahalik, Burns, and Syzdek 2007; Seidler et al. 2016; Springer and Mouzon 2011). The most influential theories of masculinity and health posit that men use unhealthy behaviors as a means of demonstrating their conformity to hegemonic masculine ideals and preserving their patriarchal privilege and status (Connell 2012, 2005; Connell and Messerschmidt 2005; Courtenay 2000). Much less work has been done on femininities and health, and in particular little research has examined

the relationship between women's health and traditional gender role beliefs that relegate women to subordinate roles in the family and society. These traditional gender role beliefs have been linked to a variety of other important outcomes in women's lives including educational attainment, age at first birth, household division of labor, and earnings (Bianchi et al. 2000; Christie-Mizell et al. 2007; Davis and Greenstein 2009; Davis and Pearce 2007; Stewart 2003). Based on these findings it is reasonable to expect that traditional gender role beliefs could influence women's health and well-being through these pathways, but also through psychosocial mechanisms like self-esteem, mastery and subjective social status which are well-known mediators of the relationship between subordinate social positions and health (Pearlin et al. 1981, 2005).

In sum, while a great deal of past research has been conducted on different topics in the area of gender inequality and health, this work has not been done within a coherent structural framework that allows for the simultaneous exploration of different levels of the gender system. Work on how macro-structural gender discrimination shapes health has been particularly scarce. This study is the first to conceptualize and measure structural sexism at different levels of the gender system and examine its relationship to physical health among women and men in the U.S. The structural sexism and health approach allows this study to consider the impact of discriminatory systems on both marginalized and dominant group members. Doing so allows us to determine whether the patterns of health effects support a conflict perspective (in which one group

benefits while the other suffers), a theory of universally harmful inequality (in which everyone suffers, although perhaps to varying degrees), or neither.

### **3.4. Methods**

Examining how structural sexism at different levels of the gender system relates to physical health among U.S. women and men requires connecting individual characteristics and health information to data reflecting the marriages and state-level environments in which individuals live. I measure macro-structural sexism using U.S. state-level administrative data compiled from a variety of sources to reflect gender inequality in political, economic, cultural and reproductive domains. I combine these data with restricted geo-code data from the National Longitudinal Study of Youth 1979 (NLSY79) to locate individuals within states to capture their exposure to structural sexism and examine how this is related to their health. I use spousal- and individual-level data from the NLSY79 to measure exposure to structural sexism at the meso- and micro-levels.

#### **3.4.1. Sample**

The NLSY79 is a nationally representative sample of individuals who were born in the years 1957 to 1964. Data were collected annually from 1979 to 1994 and biennially since 1994. The main sample includes 6,111 respondents who were age 14-22 when they were first interviewed in 1979 (U.S. Bureau of Labor Statistics 2016). The original survey also included military and supplemental samples that were dropped from the survey

prior to the study period, so only the main sample is considered in the present analysis. I use data from the 1998 through 2012 waves because this is the period during which detailed health information was collected. Respondents completed an age 40+ and an age 50+ health module on a rotating basis during the next survey year after they reached these ages. To be included in the analytic sample, respondents had to participate in the age 40 and age 50 health modules by 2012 (n=3,433). Of those eligible, 56 individuals (1.6%) were missing geo-code information and were excluded from the sample, yielding a final analytic sample of 3,377 individuals. Item missing is negligible and is therefore handled using listwise deletion.<sup>3</sup> Each individual contributes one or two person-years of information depending upon the analysis. Because meso-level structural sexism is measured in the marital dyad, analyses involving meso-level sexism use only the person years during which individuals are married. Analyses are conducted separately for men and women. Sample sizes range from 3,075 to 1,719 person-years and are noted in the tables for each analysis.

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<sup>3</sup> Missing data was less than 1% for any given outcome, and 6% or less for all independent variables except income (which was missing in 14% of cases). Listwise deletion remains an appropriate strategy in this case even given the percent missing on income for several reasons. First, sample sizes are large enough that the reduction in statistical power is unproblematic. Second, prior work has shown that listwise deletion may be less biased than standard multiple imputation or FIML methods when data is missing not at random (MNAR)-- which is often true of income if, for example, people with higher income are less likely to report (Allison 2014, 2001).

### 3.4.2. Measures

*Health Outcomes.* The relationship between structural sexism and health is assessed using three health outcomes: chronic conditions, self-rated health, and physical functioning. Chronic conditions are measured using a count of the number of major chronic conditions respondents have been told by a doctor that they have: high blood pressure/hypertension, diabetes/high blood sugar, cancer (excluding skin cancer), lung disease, heart disease/problems, stroke, psychiatric conditions, and arthritis/rheumatism (Brown, O'Rand, and Adkins 2012; Ferraro and Farmer 1999; Gorman, Read, and Krueger 2010). Self-rated health is measured on a 5-point scale from poor to excellent (Idler and Benyamini 1997). Physical functioning is measured using the physical component score portion of the SF-12, a 12-question health survey designed to provide a measure of the respondents' mental and physical health irrespective of their proclivity to use formal health services (Ware, Kosinski, and Keller 1996). Table 1 shows the health outcomes at age 40 and 50 by gender for all individuals and for married individuals. Compared to the total sample, married individuals had fewer health problems and less variance in health at both ages.

**Table 4. Health Problems by Gender at Age 40 and 50 among all Individuals and Among Married Individuals**

	Age 40				Age 50				Range
	<u>Total</u>		<u>Married</u>		<u>Total</u>		<u>Married</u>		
	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.	
WOMEN									
Chronic Conditions	0.56	0.88	0.51	0.81	1.17	1.28	1.02	1.16	[0,7]
Self-rated health	2.32	1.01	2.20	0.97	2.58	1.06	2.43	0.99	[1,5]
Physical Functioning	48.32	8.76	47.55	7.88	51.29	10.88	50.04	9.92	[33.4, 88.8]
MEN									
Chronic Conditions	0.43	0.73	0.39	0.68	0.98	1.13	0.89	1.07	[0,7]
Self-rated health	2.24	0.96	2.18	0.93	2.49	1.03	2.35	0.96	[1,5]
Physical Functioning	47.00	6.73	46.83	6.57	49.57	9.00	48.67	8.01	[32.8, 88.4]

Note: Range represents observed range across both waves. The theoretical range for physical function is 1-excellent to 100-poor.

*Macro-Level Structural Sexism.* This type of structural sexism refers to systematic gender inequality in major macro-level social institutions. Similar to other recent studies of structural discrimination (Chen et al. 2005; Kawachi et al. 1999; Lukachko, Hatzenbuehler, and Keyes 2014; Lucas 2013), the present study measures macro-structural sexism at the U.S. state-level. Table 2 describes the key indicators, all of which are designed to capture the degree to which men and women are unequal in various domains. The political and economic measures parallel those used to measure structural racism (see Lukachko et al. 2014). The political measure is the proportion of state legislature seats that are occupied by men, which previous research has shown is related

the higher infant mortality rates (Homan 2017). The three economic measures are ratios of: men's to women's labor force participation, men's to women's wages, and women's to men's poverty rates. One previous study has shown the latter two indicators to be associated with increased odds of mortality among men (Kavanagh et al. 2017). In addition to the political and economic domains, I add two additional domains that have particular relevance for structural sexism: cultural and physical/reproductive measures. For the cultural measure, I use the percentage of the state population that is composed of religious conservatives. This group includes Evangelical Protestants and Mormons (Steensland et al. 2000). The prevalence of religious conservatives is an important indicator of structural sexism because these religions relegate women to subordinate roles in the family and the church in both their ideology and practice (CBMW 2018; Chaves and Eagle 2015). The percentage of religious conservatives reflects the centrality and influence of these ideologies and practices within the state environment. Furthermore, the proportion of religious conservatives in a state has a significant relationship to conservative gender attitudes even after controlling for individuals' own religious beliefs and practices (Moore and Vanneman 2003), suggesting an important contextual effect of religious conservatism for all residents of the community. Finally, I include as a physical/reproductive equality measure the percentage of women in the state who live in a county without an abortion provider. This is an important indicator of structural sexism because feminist scholars define gender equality as freedom from

physical coercion and behavioral constraint (Chafetz 1984:5). Reproductive choice and access to a full range of reproductive health care services are generally considered to be fundamental human rights and preconditions for women’s equal citizenship and participation in social, political, and economic institutions (Borgmann and Weiss 2003; Crane and Hord 2018).

**Table 5. Measures of Macro-Structural Sexism**

Domain	Measure	Data Source
Economic	Ratio of men's to women's median usual weekly earnings of full-time wage and salary workers	Bureau of Labor Statistics
	Ratio of men's to women's labor force participation rates, age 16+	IPMUS CPS (author calculation)
	Ratio of women's to men's poverty rate (% below federal poverty line)	IPMUS CPS (author calculation)
Political	% of state legislature seats occupied by men	Institute For Women's Policy Research
Cultural	% of state population composed of religious conservatives (Evangelical Protestant or LDS)	Association of Religious Data Archives
Physical/ Reproductive	% of women who live in a county without an abortion provider	Guttmacher Institute

Values of each indicator were obtained for each U.S. state every two years during the observation period (1998-2012), with the exception of the religious conservatives and abortion access measures, which were only available in certain years (1980, 1990, 2000, 2010 for religious conservatives, and 1988, 1992, 1996, 2000, 2006, 2008, 2012 for abortion

access). For these two measures, missing time points during the observation period are filled in using state-level linear interpolation. All macro-structural sexism measures are standardized relative to the entire observation period (6,754 person years across 50 states between 1998-2012) and summed to create an index (Cronbach's alpha= .64) reflecting the overall level of macro-structural sexism across the four domains. The index has a mean of 0, a standard deviation of 3.36, and a range from -8.2 to 11.3. Appendix B discusses the index creation in detail and includes a series of supplemental analyses showing that: (1) the results are robust to alternative specifications of the index, (2) the results are not driven by any single indicator, and (3) the individual indicators are related to the health outcomes in the same (positive) direction.

*Meso-Level Structural Sexism.* This type of structural sexism refers to status and resource inequality between men and women in interactional settings. For this study, I measure meso-level structural sexism in the marital dyad, operationalized as follows:

- Logged ratio of husband's-to-wife's past year earnings (with \$1 added to all values in order to preserve the ratio format given that some individuals reported no earnings)
- Ratio of husband's-to-wife's years of education
- Husband-to-wife age ratio in years

I choose these measures because prior research indicates that these are important determinants of bargaining power within marriages. Women who are much younger, less educated and earn less than their husbands tend to have less power or status relative to their husbands (Blood and Wolfe 1960; Chang 2016; Friedberg and Webb

2006). Furthermore, evidence suggests that the age gap between spouses is related to mortality (Drefhal 2010). These measures are standardized relative to the entire sample of married person years and summed to create an index. The index has a mean of 0, a standard deviation of 1.77, and a range of -6.19 to 16.5.

*Micro-Level Structural Sexism.* This type of structural sexism refers to individually-embodied gender inequality created through the processes of socialization, internalization, identity work and construction of selves (Risman 2004). In this study, I focus on internalized gender norms as measured by a series of 4 questions about respondents' own gender role attitudes. Respondents were asked to rate their agreement with the following statements on a 4-point scale (strongly agree, agree, disagree, strongly disagree):

- A woman's place is in the home, not the office or shop
- It is much better if the man is the achiever outside the home and the woman takes care of the home and family
- Men should share the work around the house with women (reverse-coded)
- Women are much happier if they stay home and take care of children

These items are summed to create an index with a range of 4 to 20 (Cronbach's alpha = .67). Higher scores on the index indicate higher levels of individual internalized sexism.

The index has a mean of 8.44 and a standard deviation of 2.75.

*Controls.* At the state-level, I include the following time-varying characteristics as controls: (1) the percentage of the state population that is below the federal poverty line in each year, based on data from the U.S. Census Bureau (United States Census Bureau

2016a), (2) state income inequality, measured in each year using the Gini coefficient provided by (Frank 2013) (also see Frank 2014 for details), and (3) state racial composition measured by the percentage of the state population that is nonwhite in each year based on census population estimates (United States Census Bureau 2016b). I also include a dummy variable indicating whether each state is considered part of the South based on the U.S. Census Bureau's region definitions (United States Census Bureau 2015).

At the individual level, I control for respondents' age, sex, race, education (in years), household income (in thousands of dollars), marital status, parental status, and health insurance status. I also include a dummy variable indicating whether each respondent got divorced during the observation period to account for selection out of marriage based on meso- and micro-level inequality. Finally, I control for each individual's duration of residence in the state in which their state-level exposures are measured. Table 3 shows the individual-level descriptive statistics for the full sample and the married subsample. Overall, the married subsample is similar to total sample except that it has a higher mean household income, a larger proportion having children, and a larger proportion having health insurance.

**Table 6. Individual Level Descriptive Statistics (For Person-Years in Full Sample and Married Subsample)**

	Total		Married		Range
	Mean	S.D.	Mean	S.D.	
Age (years)	45.8	4.8	45.8	4.8	[39,55]
Household income (\$k)	75.1	74.9	95.2	80.5	[0,498]
Education (years)	13.5	2.5	13.7	2.6	[0,20]
Duration in state of residence (years)	9.1	2.4	9.1	2.3	[0,10]
Men	46.3%		46.9%		
Non-white	19.3%		12.8%		
Married	64.1%		100%		
Have children	80.8%		88.7%		
Have health insurance	84.9%		91.0%		
Divorced during period	10.1%		7.9%		
n (person-years)	6754		4336		

### 3.4.3. Analytic Methods

Respondents are assigned values for state-level characteristics and individual-level characteristics based on the year in which they participated in each of the two health modules. For example, someone who was 40 years old in 1998, participated in the age 40+ health module, and resided in Florida at that time would be assigned the values of macro-level variables for Florida in 1998. This person would also be assigned values of meso- and micro- level variables based on their personal and spousal characteristics in 1998. The only exception is the micro-level sexism measures which are treated as time-invariant and were measured only in 2004 for all respondents. Each respondent contributes two person-years of data, and therefore the analysis is conducted using random effects models. To adjust for potential state-level clustering of errors, I also

estimated models using cross-classified random effects for person and state. The cross-classification was used because 9% of respondents moved between time 1 and time 2. However, there was virtually no error variance at the state-level and the coefficient estimates were nearly identical to those produced using only person-level random effects. Thus, results from cross-classified models are not shown. As a robustness test I also estimated person-level fixed-effects models, but they were relatively inefficient due to very limited within-person variation in exposure to structural sexism. The fixed effects approach also did not allow for the estimation of micro-level sexism effects because the micro-level measures are time-invariant. Hausman tests indicated that random effects are appropriate<sup>4</sup>; therefore all results presented are based on models using a person-level random effects approach. Two health outcomes, self-rated health and SF-12 Physical functioning scores, are treated as continuous. The number of chronic conditions is treated as a count variable and modeled using a Poisson regression model with person-level random effects. Results shown are robust to alternative specifications of the macro- and meso- level sexism indices (see Appendix B for supplemental analyses).

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<sup>4</sup> For example,  $\chi^2(12) = 9.16$ ,  $p = .65$ , for the Hausman test among married women in models predicting chronic conditions that contained all 3 levels of sexism exposure.

### **3.5. Results**

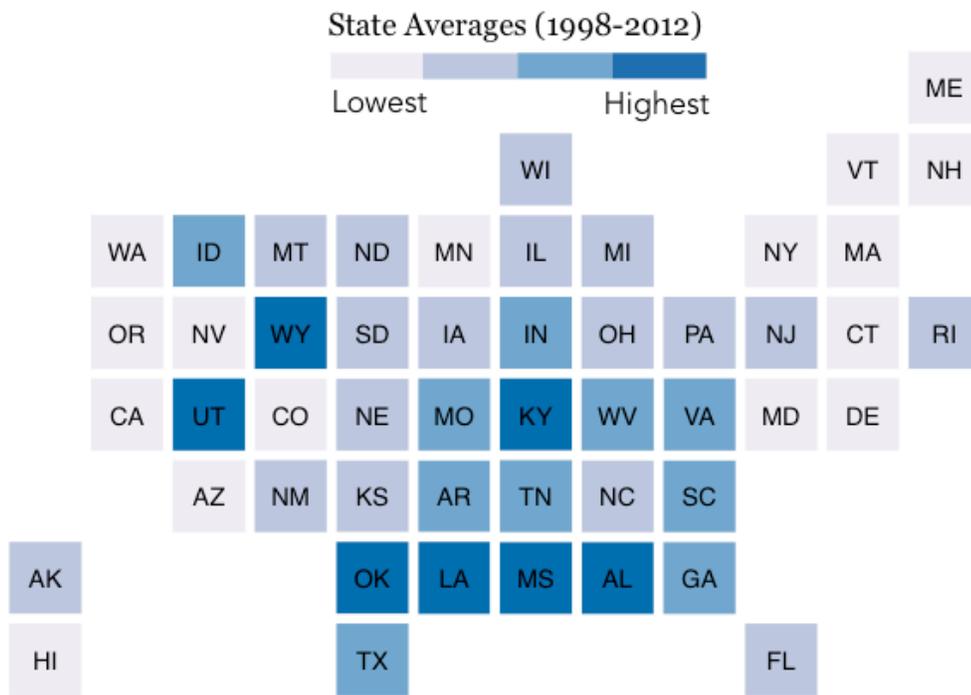
#### **3.5.1. A Descriptive Picture of Multilevel Structural Sexism in the U.S.**

*Macro-Level Structural Sexism in U.S. States.* Table 4 present the descriptive statistics for the state-level variables among the total sample of person-years. Of the six indicators of macro-structural sexism, three are ratios (wages, labor force participation, and poverty) and three are proportions (men in legislature, religious conservatives, and women lacking abortion access). For the ratio measures, a score of 1 indicates gender equality. For all three ratios, the means across all observations exceed 1, showing substantial inequality favoring men. The proportion measures also show evidence of considerable gender inequality. On average, state legislatures were composed of 77.3% men; the lowest observed proportion was 59.2% men and the highest was 95.7% men. The average proportion of the state populations composed of religious conservatives was 17.7%, and the average proportion of women living in a county without an abortion provider was 37.7%.

**Table 7. State-Level Descriptive Statistics, 1998-2012 (Among Person-Years in Full Sample)**

	Mean	S.D.	Range
Macro-Structural Sexism Index	0.00	3.36	[-8.2, 11.3]
Wage ratio (M:W)	1.27	0.07	[1.10, 1.53]
Labor force ratio (M:W)	1.21	0.05	[1.03, 1.35]
Poverty ratio (W:M)	1.25	0.12	[0.91, 2.07]
Legislature % Men	77.3%	6.4%	[59.2, 95.7]
% Religious conservatives	17.7%	10.7%	[2.3, 71.0]
% Women w/o abortion access	37.7%	22.8%	[0, 95.7]
Poverty rate	12.9%	3.1%	[4.5, 22.5]
% Population non-white	28.0%	13.1%	[2, 74]
Gini coefficient	0.601	0.041	[0.526, 0.711]
n (person-years)	6754		

Table 5 lists the states with the five highest and five lowest average levels of sexism during the study period based on each of the individual indicators and the overall macro-structural sexism index. Figure 3 is a map illustrating the geographic variation in the macro-structural sexism index scores. The map is based on each state's mean score over the time period from 1998-2012. States with the highest levels of macro-structural sexism include Utah, Wyoming, Mississippi, Louisiana, and Oklahoma, while states with the lowest average levels of macro-structural sexism include Massachusetts, California, Hawaii, Vermont and Maryland.



**Figure 5. Macro-Structural Sexism Index, State Averages 1998-2012**

**Table 8. U. S. States with the Highest and Lowest Levels of Sexism (Based on State Averages 1998-2012)**

	Macro-Sexism Index	Religious Conservatives	Lacking Abortion Access	Legislature % Male	Wage Ratio (M:W)	Poverty Ratio (W:M)	Labor Force Ratio (M:W)
Highest-1st	Utah	Utah	Wyoming	South Carolina	Wyoming	Vermont	Texas
	Wyoming	Alabama	Mississippi	Alabama	Louisiana	Indiana	Arizona
	Mississippi	Oklahoma	West Virginia	Oklahoma	Utah	Louisiana	Utah
	Louisiana	Mississippi	Arkansas	Kentucky	West Virginia	New Hampshire	California
	Oklahoma	Arkansas	South Dakota	Mississippi	Alaska	Georgia	New Jersey
Lowest-50th	Massachusetts	Connecticut	Massachusetts	Maryland	Vermont	Utah	Alaska
	California	Vermont	New York	Arizona	New York	Nebraska	Minnesota
	Hawaii	New Hampshire	Connecticut	Vermont	Maryland	Wisconsin	South Dakota
	Vermont	Massachusetts	California	Colorado	Arizona	Maine	Vermont
	Maryland	Rhode Island	Hawaii	Washington	California	California	North Dakota

*Meso-Level Structural Sexism within Marriages.* Descriptive statistics for indicators reflecting inter-spousal inequality for all person years in the married subsample are shown in Table 6. In general, husbands and wives are more closely matched in terms of education than they are in either earnings or age. On average, husbands earn \$21,000 more than their wives, yielding a mean husband-to-wife earnings ratio of 1.95. In 45% of observations, husbands earn more than twice what their wives earn. The average education difference observed between husbands and wives is zero, but in 23% of observations husbands have at least 2 years of education more than their wives. Finally, on average husbands are approximately 2 years older than their wives, yielding a mean husband-to-wife age ratio of 1.05. In 25% of couples the husband was five or more years older than the wife, whereas the reverse was true on only 5% of couples.

**Table 9. Meso-Level Descriptive Statistics (For Married Person-Years, n=4336)**

	Mean	S.D.	Range
Meso-Sexism Index	0	1.77	[-6.19, 16.52]
Husband to wife ratio of:			
Earnings (logged)	1.95	5.13	[-12.75, 12.75]
Education	1.01	0.20	[0.07, 4.00]
Age	1.05	0.12	[0.55, 2.05]
Husband earns > 2X wife	45%		
Husband education 2+ year > wife	23%		
Husband 5+ years older than wife	25%		
Wife 5+ years older than husband	5%		

Note: Both spouses in a marital dyad were not interviewed. Respondents were either a husband or a wife and report on the other spouse's characteristics.

*Micro-Level Internalized Sexism.* To what extent do men in women who were in their mid-forties in 2004 espouse an ideology of traditional gender roles? For illustrative purposes, I describe two individual indicators and then provide further detail about the index scores. Roughly 10% of the sample agreed or strongly agreed that “A woman’s place is in the home, not the office or the shop,” and 25% agreed/strongly agreed that “It is much better if the man is the achiever outside the home and the woman takes care of the home and family.” Table 7 shows the descriptive statistics for the micro-level sexism index by gender among the total sample and married subsample. Women exhibit lower internalized sexism levels than men and this difference is statistically significant at the  $p < .001$  level. Index scores among the married subsample are higher than the total sample.

**Table 10. Micro-Level Sexism Index Scores by Gender (For Person-Years in Full Sample and Married Subsample)**

Micro-Sexism Index	Total			Married		
	Total	Women	Men	Total	Women	Men
Mean	8.44	8.03	8.91	8.47	8.08	8.90
S.D.	2.75	2.79	2.63	2.77	2.79	2.62

Notes: Differences between women and men are significant at the  $p < .001$  level. There are no significant difference between the total sample and the married subsample.

### **3.5.2. The Relationship Between Structural Sexism and Health at Midlife**

Table 8 presents coefficients for the effects of structural sexism on women’s physical health. For each of the three health outcomes results are shown from two

models. Model 1 includes only macro-level structural sexism and controls. Model 2 includes all available levels of structural sexism exposure (macro- and micro- only for the full sample, and all three levels for the married subsample), plus controls. The results in Table 8 show that higher macro-structural sexism is associated with more chronic conditions, worse self-rated health and worse physical functioning for women. The effects tend to be slightly larger among the married subsample. For meso-level sexism (measured only within marriages), there are statistically significant effects on the number of chronic conditions and physical functioning. Micro-level internalized sexism was associated with more chronic conditions in the full sample, but did not have any independent effects net of the other levels of structural sexism exposure on the physical health of married women. This pattern may reflect a particularly harmful effect of internalized sexism among unmarried women who may feel that their marital status is preventing them from fulfilling their ideal of women's roles. The effects of macro-level sexism remain relatively stable when sexism at other levels is added to the model (Model 2), suggesting that the macro-, meso-, and micro levels operate relatively independently. One exception is the physical functioning among married women, for which the macro-level effects are mediated by those at the meso-level.

**Table 11. Coefficient Estimates For Structural Sexism From Random Effects Models Predicting Women's Physical Health Outcomes**

Structural Sexism Exposure	Chronic Conditions		Poor SRH		Poor Phy. Functioning	
	1	2	1	2	1	2
Among Full Sample						
Macro-level	0.027** (0.010)	0.023* (0.010)	0.015* (0.007)	0.014+ (0.007)	0.197** (0.073)	0.169* (0.076)
Micro-Level		0.022* (0.010)		0.012 (0.008)		0.081 (0.083)
n	3075	2891	3074	2890	3058	2875
Among Married Subsample						
Macro-level	0.045*** (0.013)	0.040** (0.013)	0.021* (0.009)	0.019* (0.009)	0.190* (0.084)	0.123 (0.084)
Meso-level		0.052* (0.021)		0.023 (0.016)		0.686*** (0.171)
Micro-level		0.005 (0.014)		0.007 (0.010)		-0.001 (0.092)
n	1932	1742	1932	1742	1921	1731

Notes: Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region) which are not shown.

The coefficients for the effects of structural sexism on men's physical health are shown in Table 9. The results show that macro-level structural sexism is harmful for men's physical health as well as women's. However, the effects are much more pronounced among married men. Among the married subsample, greater macro-level structural sexism is associated with more chronic conditions, and worse physical functioning; the effect on self-rated health is marginally significant. Among the full

sample of men, macro-level sexism only has significant effects for physical functioning. Interestingly, meso-level sexism affects men's health in the opposite direction. Rather than being harmful to men's health, greater meso-level sexism is associated with *better* self-rated health and better physical functioning among men. Micro-level internalized sexism is not associated with men's physical health. Comparisons between Models 1 and 2 again suggest that sexism exposures at the macro- and meso- levels of the gender system have independent effects on health.

**Table 12. Coefficient Estimates for Structural Sexism Exposure from Random Effects Models Predicting Men's Physical Health Outcomes**

Structural Sexism Exposure	Chronic Conditions		Poor SRH		Poor Phy. Functioning	
	1	2	1	2	1	2
Among Full Sample						
Macro-level	0.013 (0.012)	0.015 (0.012)	0.012 (0.008)	0.013 (0.008)	0.145* (0.060)	0.158* (0.063)
Micro-Level		-0.009 (0.014)		-0.002 (0.009)		0.000 (0.069)
n	2718	2517	2722	2521	2708	2508
Among Married Subsample						
Macro-level	0.034* (0.016)	0.033* (0.017)	0.017+ (0.009)	0.016+ (0.010)	0.228** (0.069)	0.253*** (0.073)
Meso-level		-0.033 (0.024)		-0.033* (0.014)		-0.429** (0.152)
Micro-level		-0.007 (0.018)		-0.003 (0.010)		-0.039 (0.081)
n	1793	1627	1795	1628	1788	1621

Notes: Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region) which are not shown.

Although sexism exposures at different levels have independent effects, it is possible that exposures at one level could exacerbate or buffer the effects of exposures at another level. For example, perhaps macro-level sexism is less harmful for women in more egalitarian marriages. However, I did not find evidence to support this hypothesis. I tested for cross-level interactions between macro-, meso-, and micro-level sexism among both men and women and found no statistically significant interactions. This means the effect at each level are additive. On the basis of intersectionality theory and research (cf. Brown et al. 2016), I also tested for interactions between an individual's race and their sexism exposures among men and women separately, and again found no statistically significant interaction effects.

To summarize the results, Figure 4 illustrates the predicted health problems across differing degrees of exposure to structural sexism at each level of the gender system among married women and men. The stars indicate the effects that are significantly different from zero in the single gender models shown in Tables 8 and 9 (a single star is used to indicate all levels of statistical significance of at least  $p < .05$ ).<sup>5</sup> Macro-structural sexism is associated with worse health among both men and women,

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<sup>5</sup> To generate these predicted values men and women were included in a single model with an interaction between gender category and sexism exposure. This strategy makes it possible to hold all other variables constant at their overall mean, but it assumes that the effects of these other variables are equal across gender category – an assumption that was not made in the separate models by gender for which the results are presented in Tables 8 and 9. This strategy also makes it possible to test whether the effects of sexism differ for men and women. Results indicate that at the macro- and micro- levels, the effects are not significantly different for men and women, while at the meso-level they are significantly different and, in fact, are opposite in direction.

supporting a theory of universally harmful inequality. Based on model predictions, 45 year old women with levels of macro-structural sexism exposure 2 standard deviations below the mean have an average of 0.5 chronic conditions, while similar women at two standard deviations above the mean exposure have almost twice as many chronic conditions. This difference is roughly equivalent to the effect of being 7 years older. Although the effect sizes of macro-structural sexism among men are slightly smaller than those observed among women, the difference is not statistically significant.

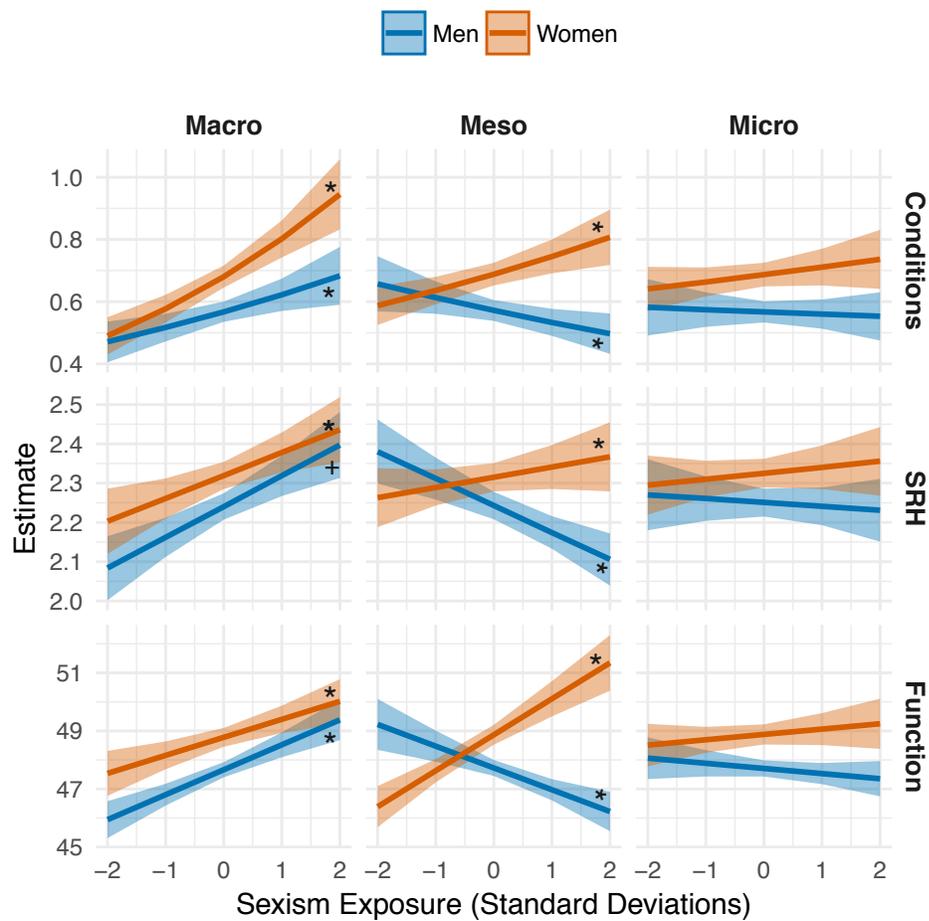


Figure 6. Predicted Values of Health Outcomes Given Structural Sexism Exposure

At the meso-level, however, the patterns in effects across gender category are strikingly different. Greater structural sexism between husbands and wives is associated with more physical health problems among women, but fewer physical health problems among men. This pattern supports a theory of zero-sum gender conflict in which men reap health benefits from greater dominance or status relative to their wives. At the micro-level, internalized sexism in the form of adherence to traditional gender role ideology was not associated with physical health outcomes among either men or women and did not mediate or moderate the effects of exposures at other levels.

### **3.6. Discussion and Conclusions**

Rather than focus on gender differences in health outcomes or on the harmful effects of perceived discrimination and harassment on women, I have advanced a *structural sexism* approach to the study of gender inequality and health that is more consistent with contemporary theoretical understandings of gender as a multilevel social system. This approach directs our attention beyond individual actors to the question of how the inequitable gendered distribution of power, status, and resources that characterizes a society's gender system influences the health of its members. This study is the first to conceptualize and measure structural sexism at different levels of the gender system and to examine its relationship to physical health among both women and men in the U.S. The findings show that structural sexism has important health consequences for both women and men in midlife.

For women, the results show more physical health problems are associated with greater exposure to structural sexism at both the macro- and meso- levels, which aligns with logical expectations given their disadvantaged social position. Existing health disparities work offers substantial insight into the process of how social inequality “gets under the skin.” For example, Nancy Krieger's (2014, 2001) ecosocial theory describes the processes through which discriminatory social systems undermine the health of marginalized groups by structuring their living conditions and the exercise of their civil, political, social, economic and cultural rights in ways that create economic and social deprivation, toxic/hazardous living conditions, socially inflicted trauma, and inadequate healthcare. The ecosocial model in combination with other health disparities research suggests a variety of specific pathways through which structural sexism may harm women's health including: reduced access to material resources, goods, and services; exposure to violence, harassment or unsafe working conditions; perceived discrimination; low subjective social status; increased stress; reduced psychosocial resources such as self-esteem, mastery, sense of control, coping resources and social support; and inadequate health care quality and/or access (Aizer 2010; Link and Phelan 1995; Marmot 2005; Pascoe and Richman 2009; Pearlin et al. 2005; Yang, Schorpp, and Harris 2014). All of these factors may contribute to the positive association between structural sexism and health problems among women, with the material resources and exposures possibly playing a larger role at the macro-structural level and the psychosocial resources and stress playing a larger role at the meso- and micro- levels.

Among men, however, understanding the health impacts of structural sexism is less straightforward. Conflict theories in general, and the ecosocial model in particular, suggest that dominant groups benefit at the expense of their subordinates (Collins 1975; Krieger 2014:74). As the dominant group, men are expected to experience increasing levels of power, resources and status -- and thus a corresponding health benefit -- with increasing degrees of structural sexism. The results at the meso-level support this type of gender conflict perspective, with greater degrees of inter-spousal structural sexism being associated with fewer health problems among men and more health problems among women. The bounded nature of a marital dyad means that resources and responsibilities are more likely to be zero-sum and the spouse with the higher status position stands to benefit more directly from inequality in ways that may not hold true for larger aggregate-level inequalities.

In fact, patterns in the effects of structural sexism on men's health at the macro-level do not support a zero-sum gender conflict perspective. At the macro-level, greater exposure to structural sexism is related to worse health among *both* women and men, showing a pattern of universally harmful inequality. This pattern is consistent with broader theoretical perspectives on structural inequalities and health which argue that inequality harms everyone because it damages social relationships, increases competition for dominance, undermines the social fabric, and makes the entire society less safe, less productive, and less healthy (Lucas 2013; Wilkinson 2005; Wilkinson and Pickett 2011).

The pattern of universal harm is also consistent with theory and research specifically focused on gender inequality and health. Research in the developing world has shown that gender equity is vital for development, and poverty reduction and improvements in population health (World Bank 2003, 2011). Studies suggest that when women are empowered they influence social policy in ways that promote education, health care and social programs and expenditures that improve population health (Boehmer and Williamson 1996; Bolzendahl and Brooks 2007; Little et al. 2001; Miller 2008). The observed pattern of universal harm supports gender theories which posit that patriarchal social systems foster toxic constructions of masculinity that shape both institutions (Acker 1992) and cultural norms in way that harm men's health (Connell 2005, 2012; Courtenay 2000).

The lack of a statistically significant relationship between physical health and internalized gender ideology at the micro-level is surprising given prior work on masculinity and health (Mahalik et al. 2007; Seidler et al. 2016; Springer and Mouzon 2011) and it implies that the effects of hegemonic masculinity on health have an important institutional component that is not fully captured by individual beliefs and behaviors. The null findings at the micro-level may also indicate that the specific gender role beliefs examined here are: (1) less important for health than other aspects of masculinity and gender norms, and/or (2) more closely associated with other non-physical health outcomes not considered here such as mental health and life satisfaction.

This study represents an early attempt to characterize the relationship between structural sexism and health, similar to emerging work on the health consequences of structural racism (Bailey et al. 2017). As such, future research is needed to refine and extend both the theory and measurement of structural sexism in several ways. First, future research should develop additional measures of structural sexism across a variety of domains, with particular attention to the meso-level. The present study examined meso-level structural sexism only within the context of heterosexual marriages<sup>1</sup>. Although heterosexual marriage is a key site for the reproduction of gender inequality, future research should also investigate structural sexism in a variety of other interactional settings including the neighborhood and the workplace. Second, future research should examine the effects of structural sexism on other health outcomes (including mental health and mortality) and with longitudinal data over a longer time period to gain a more complete picture of how structural sexism shapes trajectories of health decline in later life. Third, the present study did not contain measures of perceived interpersonal gender discrimination or sexual harassment. Investigating the relationship between structural sexism and this more overt type of gender-based mistreatment and assessing the role such mistreatment plays in the structural sexism-health relationship are vital next steps for future research. Finally, a critical task for future research will be to move beyond the gender binary theoretically and empirically.

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<sup>1</sup> Data limitations did not allow for the examination of same-sex marriages in the present study, but recent work has shown that gendered processes related to health behaviors operate differently for men and women in gay, lesbian, and straight relationships (Reczek 2012; Reczek and Umberson 2012).

This will entail a twofold process of working to better understand the impact of discriminatory gender systems on transgender individuals and those with non-binary gender identities, as well as grappling with how measurement of structural sexism relates to the gender binary. Although the gender system in American society is still fundamentally organized around heterosexual relationships and the man-woman gender binary, an increasing number of individuals are identifying as transgender and non-binary (Meerwijk and Sevelius 2017). To the extent that institutions adjust to this reality and shift the gendered distribution of power and resources, new ways of measuring structural sexism which take this into account may become increasingly necessary.

The results of this study demonstrate the promise of a structural sexism approach to gender inequality and health. The fact that macro- and meso-level results support different theories regarding the effects of inequality on the health of dominant group members highlights the importance of a *multilevel* systems approach. Overall, the findings suggest that systematic gender inequality in the U.S. is both a human rights issue and a public health problem. The health of our entire society is likely to be undermined by the systematic exclusion of women from resources and power within major social institutions.

## 4. Conclusion

In this dissertation, I developed a theoretical framework and measures for the concept of *structural sexism* and examined its relationship to health in the United States. Consistent with contemporary theories of gender as a multilevel social system, I conceptualized and measured structural sexism as *systematic gender inequality in power and resources* at the macro-level (U.S. state), meso-level (marital dyad), and micro-level (individual). Through a series of quantitative analyses, I examined how various measures of structural sexism affect the health of men, women, and infants in the U.S.

In chapter 2, I began by focusing on macro-structural sexism in the political domain. I used data compiled from several administrative sources across a twenty-two-year period, to show that women's political representation in U.S. state legislatures is as strong — or stronger — a predictor of state infant mortality rates than even the state's poverty level. This relationship holds both between states with varied levels of political gender inequality in a given year, and within the majority of states as women's representation changes over time. To quantify the population-level costs, I estimated that women's continued underrepresentation in legislative office is associated with nearly 3,500 excess infant deaths per year in the U.S. compared to what would be expected under conditions of gender parity.

In chapter 3, I laid out a theoretical framework for structural sexism. I developed measures of structural sexism at the macro-, meso-, and micro-levels of the gender system and examined its consequences for the health of women and men at midlife. I

measured macro-level structural sexism at the U.S. state-level using indicators of inequality in political, economic, cultural and reproductive domains. Using restricted geo-coded data I located individuals within states to capture their exposure to macro-structural sexism. I also incorporated individual- and spousal-level data in order to measure exposure to structural sexism at the meso- and micro-levels. Results showed that among women exposure to more sexism at the macro- and meso-levels is associated with more chronic conditions, worse self-rated health, and worse physical functioning. Among men, macro-level structural sexism is also associated with worse health. However, at the meso-level greater structural sexism is associated with better health among men. At the micro-level, internalized sexism is not related to health among either women or men. These results highlight the importance of a multilevel approach.

As a whole, this work demonstrates that structural sexism takes a substantial toll on population health in the U.S. The findings also illustrate that effects of structural sexism on health depend both on an individual's position within the gender system and the level at which sexism exposures are measured. Future research should explore the mechanisms through which structural sexism impacts health, and situate structural sexism with respect to overlapping systems of economic and racial oppression. This line of research will continue to make vital contributions to our understanding of how structural inequalities shape health across the life course.

## **Appendix A**

This appendix includes the supplementary discussion, tables and figures for Chapter 2. Figure 7 illustrates the bivariate relationships between IMRs and key control variables. The first panel shows that the average state IMR decreases over time. The second and third panels demonstrate that lower infant mortality is associated with higher levels of median income and educational attainment. The fourth panel shows a weak to nonexistent relationship between state racial composition and IMR.

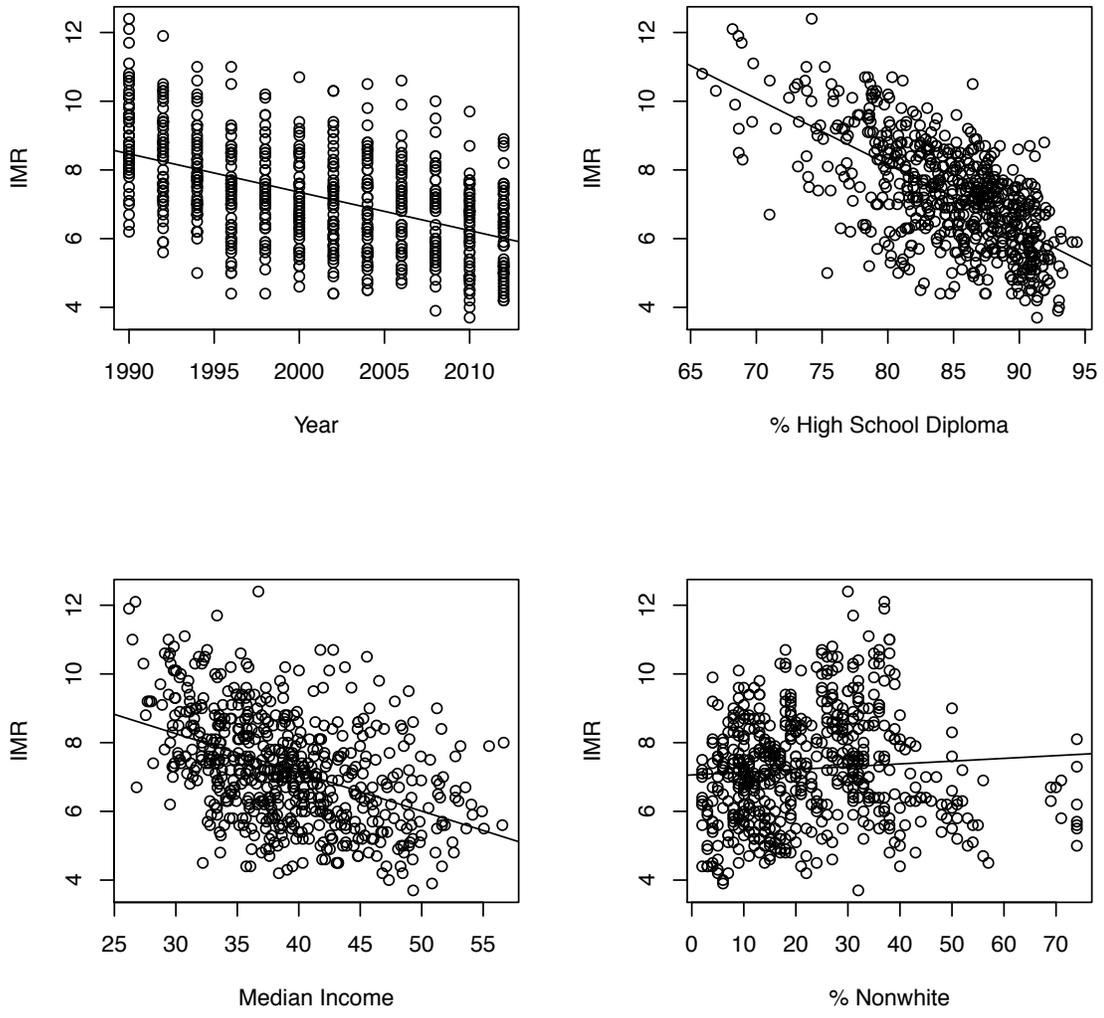


Figure 7. IMR by Selected Controls, U.S. States, 1990-2012 (n=600)

Table 13 shows results from supplemental analyses that are similar to model R3 but also include political party variables. Model 1 includes a series of dummy variables indicating partisan control of each state legislature as either Republican majority, democratic majority, or split, in each year. These additional partisan control variables do not have a statistically significant relationship to infant mortality, and the model results do not substantively differ from those presented in Table 3. Furthermore, the Wald test indicates that adding these additional controls did not create a statistically significant improvement in model fit (Wald  $\chi^2(2)=0.84$ ,  $P=0.657$ ). These results suggest that the relationship between political gender inequality and infant mortality cannot be attributed solely to the partisan composition of legislatures.

Model 2 includes a control for the political party affiliation of the sitting U.S. President (0=Democrat, 1=Republican). The results show that the main findings of this study are robust. The coefficient for women's representation decreases slightly in magnitude and significance but remains negative and statistically significant. The results also confirm Rodriguez et al.'s (2013) finding Republican administrations are associated with higher IMRs, suggesting that national political processes may impact infant mortality.

**Table 13. Results of Longitudinal Regression Model Predicting U.S. State Infant Mortality Rates Including Political Party Variables, 1990-2012**

	Model 1	Model 2
Legislature % Women	-0.036** (0.01)	-0.029* (0.01)
Year	-0.092*** (0.01)	-0.094*** (0.01)
%	0.028 (0.03)	0.031 (0.03)
Southern Region (ref: non-south)	1.311*** (0.29)	1.337*** (0.28)
% HS Degree or Higher	-0.021 (0.02)	-0.011 (0.02)
Median Income	0.002 (0.02)	-0.012 (0.02)
Gini	1.174 (2.39)	0.351 (2.33)
% Nonwhite	-0.003 (0.01)	-0.001 (0.01)
Partisan Control of Legislature (ref: split)		
Republican	-0.073 (0.08)	-0.045 (0.07)
Democrat	-0.038 (0.09)	-0.056 (0.09)
Republican President (ref: Democratic)		0.337*** (0.06)
Constant	193.081*** (25.81)	195.828*** (26.19)
n	588	588
Model R <sup>2</sup> : Overall	0.576	0.595
Between-state	0.565	0.581
Within-state	0.593	0.617

\* p<.05, \*\* p<.01, \*\*\*p<.001, two-tailed tests. Robust clustered standard errors are included in parentheses.

Note: All observations for Nebraska excluded due to non-partisan

Table 14 demonstrates that the results in Table 3 are robust to several alternative specifications of the lag period between women's political representation and infant mortality. Models show results with no lag (equivalent to models R4 and F4 in Table 3), and lags of 2, 4 and 6 years. Estimates remain stable and statistically significant for lags between 0 and 4 years. In models with a 6-year lag, estimates decrease somewhat and fall just below the threshold of statistical significance. This may be a function of declining sample sizes, a waning relationship, or both.

**Table 14. Results of Longitudinal Regression Models Predicting U.S. State Infant Mortality Rates with Alternative Lag Periods, 1990-2012**

	Random-Effects				Fixed-Effects			
	Lag 0	Lag 2	Lag 4	Lag 6	Lag 0	Lag 2	Lag 4	Lag 6
Legislature % Women	-0.032* (0.01)	-0.031* (0.01)	-0.030* (0.01)	-0.021 (0.01)	-0.025** (0.01)	-0.026** (0.01)	-0.025* (0.01)	-0.013 (0.01)
Year	-0.096*** (0.01)	-0.086*** (0.01)	-0.078*** (0.01)	-0.066*** (0.02)	-0.089*** (0.01)	-0.088*** (0.01)	-0.084*** (0.02)	-0.063*** (0.02)
% Poverty	0.028 (0.03)	0.013 (0.03)	-0.006 (0.02)	-0.017 (0.03)	0.034 (0.02)	0.018 (0.02)	-0.001 (0.02)	-0.005 (0.02)
Southern Region (ref: non-	1.351*** (0.28)	1.414*** (0.28)	1.555*** (0.27)	1.718*** (0.26)				
% HS Degree or	-0.019 (0.02)	0.001 (0.02)	0.015 (0.02)	0.000 (0.02)	-0.026 (0.02)	0.001 (0.02)	0.021 (0.02)	0.004 (0.02)
Median Household	0.002 (0.02)	-0.013 (0.02)	-0.012 (0.02)	-0.001 (0.02)	0.017 (0.02)	0.002 (0.02)	0.009 (0.02)	0.037 (0.02)
Gini	1.314 (2.29)	0.150 (1.77)	-0.668 (1.54)	-0.537 (1.39)	1.876 (1.37)	0.542 (1.35)	-0.464 (1.38)	-0.109 (1.39)
% Nonwhite	-0.003 (0.01)	0.003 (0.01)	0.005 (0.01)	-0.000 (0.01)	-0.031 (0.02)	-0.006 (0.02)	0.001 (0.03)	-0.022 (0.03)
Constant	200.937** (24.91)	179.403*** (24.71)	162.839** (27.25)	138.684** (31.45)	187.523** (25.61)	183.408** (26.61)	173.449** (28.61)	132.811** (30.39)
n	600	550	500	450	600	550	500	450

\* p<.05, \*\* p<.01, \*\*\*p<.001, two-tailed tests.

Note: Standard errors included in parentheses.

Table 15 shows the variance inflation factor for each variable in the full model (R4). The VIF estimates are based on standard OLS models (using the full sample of 600 state years) because in random or fixed-effects models the standard errors are not calculated in the same way as in linear models and therefore the VIFs do not technically apply to these models. Nonetheless, all estimated VIFs are below 5 indicating that multicollinearity is not a major concern. In particular the VIF for legislature % women is only 1.70. The VIFs for % high school graduates and % poverty are somewhat higher (although still below thresholds considered cause for concern) and may result in slightly less stable estimates and inflated standard errors for these two control variables.

**Table 15. Variance Inflation Factors for All Variables in Full Model (R4)**

	VIF
Legislature % Women	1.70
Year	2.68
% HS Degree or Higher	4.20
% Poverty	4.22
South	1.92
Median Income	3.52
Gini	1.48
% Nonwhite	2.03
Mean	2.72

Table 16 and 17 show additional fixed-effects and random-effects models that highlight the role of educational attainment and clarify why the coefficients for % high school degree or higher are nonsignificant in fully adjusted models in the main analysis (Models R4 and F4 in Table 3). Fixed-effects results (16) show that when only educational attainment and legislature % women are included in a fixed effects model predicting IMRs in the U.S., both are statistically significant (Model 1). When year is added to the model, the effects of year and legislature % women are statistically significant, but the effect of education becomes nonsignificant (Model 2). When legislature % women is excluded from the model, the effect of education is still nonsignificant (Model 3). Therefore, it appears that year functions as a proxy for educational attainment, in addition to other unmeasured medical and technological advances which may have reduced IMRs in the U.S. over time.

**Table 16. Alternative Specifications of Fixed-Effects Regression Models Predicting Infant Mortality Rates, United States, 1990-2012**

	Model 1	Model 2	Model 3
% HS Degree or	-0.16***	-0.02	-0.03
Legislature % Women	-0.06***	-0.03*	
Year		-0.09***	-0.10***

\* p<.05, \*\* p<.01, \*\*\*p<.001, two-tailed tests. Robust clustered standard errors are included in parentheses.

In the random-effects framework (17), the coefficient for %HS degree retains significance after the year is included in the model, but the coefficient shrinks to roughly a quarter of its original size. The coefficient for %HS degree becomes non significant when the dummy variable for southern region is introduced. In sum, the coefficient for educational attainment depends on which other controls are included in the model. But more importantly the coefficient of interest, legislature % women, remains statistically significant and relatively stable regardless of which covariates are included in or excluded from the models.

**Table 17. Alternative Specifications of Random-Effects Regression Models Predicting Infant Mortality Rates, United States, 1990-2012**

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
% HS Degree or Higher	-0.192*** (0.01)	-0.052** (0.02)	-0.042* (0.02)	-0.035* (0.02)	-0.018 (0.02)	
Year		-0.088*** (0.01)	-0.095*** (0.01)	-0.087*** (0.01)	-0.098*** (0.01)	-0.102*** (0.01)
%Poverty			0.030 (0.02)	0.026 (0.03)	0.026 (0.02)	0.032* (0.02)
Median Household Income			-0.010 (0.02)	-0.005 (0.02)	0.000 (0.02)	
Gini			1.540 (2.16)	1.276 (2.43)	1.245 (2.29)	
Legislature % Women				-0.040** (0.01)	-0.033* (0.01)	-0.034** (0.01)
Southern Region (ref: non-south)					1.327*** (0.29)	1.401*** (0.28)
Constant	23.520*** (1.19)	187.529*** (18.38)	200.638*** (21.29)	184.207*** (24.08)	203.116*** (24.28)	211.922*** (13.83)
n		600	600	600	600	600
Model R <sup>2</sup> : Overall	0.417	0.353	0.394	0.484	0.577	0.579
Between-state	0.379	0.379	0.389	0.497	0.566	0.571
Within-state	0.47	0.583	0.585	0.591	0.595	0.592

\* p<.05, \*\* p<.01, \*\*\*p<.001, two-tailed tests. Robust clustered standard errors are included in parentheses.

## Appendix B

Below I discuss measurement issues pertaining to the creation of the structural sexism indices in Chapter 3. At the macro-level, the main considerations were how to combine the six items to create a summary measure of state-level structural sexism across political, economic and cultural domains, and whether the results are robust to alternative approaches. At the meso-level the primary consideration was how to measure the amount of inequality between husbands and wives in terms of age, education, and income, and again whether the results are robust to alternative approaches. I do not discuss measurement at the micro-level because the items were standard gender role beliefs, all measured on the same scale, and thus could be combined as a simple sum.

*Macro-Level Sexism Index.* To create the index used in the main analyses, each of the six state-level indicators was standardized relative to the entire sample of person-year observations (representing the state-level conditions to which respondents were exposed at ages 40 and 50 between 1998-2012) and then summed to create the index. Tables 18 & 19 compare the coefficient estimates for married women and men produced using this approach to estimates produced using two alternative approaches. For alternative 1, I logged the three ratio measures (labor force participation ratio, wage ratio, and poverty ratio) and then subtracted 50% from the proportion of legislators that are men in order to center this measure around zero because 50% is the theoretical

equality point. Then the measures were standardized relative to the entire sample of person-year observations and summed to create an index. Alternative 2 was the same as the original index except that instead of standardizing each indicator across all person-year observations, indicators were standardized within each of the two person-year observations (because of the data structure this means that the first observation for each person was exposure at age 40, and was between 1998-2004 and the second observation was exposure to sexism at age 50 measured between 2006-2012). As tables 18 and 19 show, these alternative approaches did not produce meaningfully different estimates.

**Table 18. Comparison of Macro-Sexism Results Using Alternative Specifications of Index Among Married Women**

	Conditions	Poor SRH	Poor Functioning
Macro-Sexism Index	0.045*** (0.013)	0.021* (0.009)	0.190* (0.084)
Alternative 1	0.045*** (0.013)	0.020* (0.009)	0.183* (0.085)
Alternative 2	0.044*** (0.013)	0.020* (0.009)	0.188* (0.082)
n	1932	1932	1921

Notes: Among married women subsample. Coefficient estimates are from random effects models leaving items out of the index one at time. Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include meso- and micro-level sexism indices.

**Table 19. Comparison of Macro-Sexism Results Using Alternative Specifications of Index Among Married Men**

	Conditions	Poor SRH	Poor Functioning
Macro-Sexism Index	0.034* (0.016)	0.017+ (0.009)	0.228** (0.069)
Alternative 1	0.033* (0.016)	0.018* (0.009)	0.230*** (0.069)
Alternative 2	0.034* (0.015)	0.016+ (0.009)	0.237*** (0.068)
n	1932	1932	1921

Notes: Among married men subsample. Coefficient estimates are from random effects models leaving items out of the index one at a time. Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include meso- and micro-level sexism indices.

Tables 20 and 21 present coefficients for married women and men from a series of leave-one-out models estimated as an additional robustness test. These models were identical to the original analyses except each model left out a single item from the macro-structural sexism index. The results in Tables 19 and 20 indicate that the main results reported in this study are not driven by any single indicator of macro-structural sexism.

**Table 20. Leave-One-Out Replication of Macro-Sexism Results Among Married Women**

Item Left Out	Conditions	Poor SRH	Poor Functioning
Labor Force Ratio	0.050*** (0.015)	0.022* (0.010)	0.198* (0.095)
Poverty Ratio	0.067*** (0.016)	0.034** (0.011)	0.412*** (0.104)
Male legislators	0.044** (0.015)	0.023* (0.011)	0.189+ (0.101)
Limited Abortion Access	0.044** (0.014)	0.020* (0.010)	0.148 (0.092)
Religious Conservatives	0.038** (0.014)	0.021* (0.010)	0.155+ (0.091)
Wage Ratio	0.051** (0.016)	0.020+ (0.011)	0.170+ (0.100)
Complete Index	0.045*** (0.013)	0.021* (0.009)	0.191* (0.084)
n	1932	1932	1921

Notes: Among married women subsample. Coefficient estimates are from random effects models leaving items out of the index one at time. Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include meso- and micro-level sexism indices.

**Table 21. Leave-One-Out Replication of Macro-Sexism Results Among Married Men**

Item Left Out	Conditions	Poor SRH	Poor Functioning
Labor Force Ratio	0.032* (0.016)	0.010 (0.010)	0.220** (0.076)
Poverty Ratio	0.042* (0.019)	0.014 (0.012)	0.224** (0.080)
Male legislators	0.036* (0.019)	0.024* (0.011)	0.218** (0.083)
Limited Abortion Access	0.034* (0.017)	0.018+ (0.010)	0.229** (0.074)
Religious Conservatives	0.035* (0.017)	0.021* (0.010)	0.276*** (0.077)
Wage Ratio	0.038* (0.018)	0.022* (0.010)	0.290*** (0.082)
Complete Index	0.034* (0.016)	0.017+ (0.009)	0.228** (0.069)
n	1793	1795	1788

Notes: Among married men subsample. Coefficient estimates are from random effects models leaving items out of the index one at time. Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include meso- and micro-level sexism indices.

Finally, Tables 22 and 23 show the results for each individual item (instead of the index) as a predictor of health outcomes. These tables again show that no single indicator is driving the index results. The tables also show that the indicators are all related to health in the same direction: greater sexism is associate with worse health. Only five out of the forty eight coefficients estimated are negative and these are not statistically significant.

**Table 22. Individual Macro-Sexism Indicators Among Married Women**

	Conditions	Poor SRH	Poor Functioning
Macro-Sexism Index	0.045*** (0.013)	0.021* (0.009)	0.191* (0.084)
Male legislators	1.823** (0.596)	0.599 (0.443)	7.300+ (3.868)
Religious Conservatives	1.547** (0.488)	0.527 (0.387)	9.409* (4.149)
Limited Abortion Access	0.691** (0.261)	0.428* (0.192)	6.483*** (1.816)
Poverty Ratio	0.039 (0.240)	-0.033 (0.164)	-2.995+ (1.601)
Labor Force Ratio	0.526 (0.656)	0.516 (0.541)	4.179 (5.078)
Wage Ratio	1.026* (0.504)	0.883* (0.388)	8.351* (3.790)
n	1932	1932	1921

Notes: Coefficient estimates are from random effects models using married women subsample. Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include meso- and micro-level sexism indices.

**Table 23. Individual Macro-Sexism Indicators Among Married Men**

	Conditions	Poor SRH	Poor Functioning
Macro-Sexism Index	0.034* (0.016)	0.017+ (0.009)	0.228** (0.069)
Male legislators	1.020 (0.661)	-0.019 (0.394)	9.419** (3.007)
Religious Conservatives	0.502 (0.565)	-0.187 (0.340)	-1.516 (2.733)
Limited Abortion Access	0.301 (0.321)	0.093 (0.189)	2.256 (1.394)
Poverty Ratio	0.226 (0.284)	0.288+ (0.159)	3.128* (1.222)
Labor Force Ratio	1.031 (0.818)	1.398** (0.516)	6.849+ (4.053)
Wage Ratio	0.715 (0.561)	0.117 (0.394)	1.838 (2.832)
n	1793	1795	1788

Notes: Coefficient estimates are from random effects models using married men subsample. Robust standard errors are in parentheses. + p<.10 \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include meso- and micro-level sexism indices.

*Meso-Level Sexism Index.* For the meso-level sexism index, the key issue is how to best operationalize power and resources *inequality* between husbands and wives in terms of education, income, and age, while holding constant the overall amount of resources available that can be leveraged to maintain health. In the main analysis, I used straightforward husband-to-wife ratios of education, income, and years of age, where \$1 is added to income (to maintain ratio format in cases where \$0 income is reported) and the income ratio is logged (to reduce skewness). Then, consistent with the primary

approach used at the macro level, these ratios were standardized across all person-years in the married subsample and summed to create an index. Tables 23 and 24 compare the coefficient estimates for the effects of meso-level structural sexism on the health of women and men using this measurement strategy and four alternative strategies. In all of these models, to account for the overall amount of resources available I control for transferrable resources at the household level (i.e., household income) and non-transferrable resources at the individual level (i.e., age and education).

Alternative 1, is the same as the original index, except without logging the wage ratio. Alternative 2, measures inequality in each type of resource as the proportion of the couple's total that is attributable to the husband. For example, inequality in income would be calculated as: husband's income / (husband's income + wife's income). This was repeated for years of education and age, and then standardized and summed to create the index. Alternative 3, is the same as alternative 2 except that the three proportion indicators were not standardized before being summed. Finally, alternative 4 uses three dichotomous measures because inequality in education, income, and age may not reflect interspousal power/resource inequality in a linear way. For example, it is unclear whether a marriage in which a husband earns \$450K more than his wife is more unequal than a marriage in which a husband earns \$350K more than his wife. Therefore, I constructed 3 dichotomous measures indicting whether there was a meaningful level of inequality in each resources favoring the husband. These were coded as 1 if the

husband had 2+ years of education more than the wife (0 if not), 1 if the husband's earnings were 1.5x his wife's or more (0 if not), and 1 if the husband was 5+ years older than his wife. The three indicators were then summed to create an index with a score of 0 to 3. One of the drawbacks to this approach is that the cut-off points are somewhat arbitrary. However, I tested other cut-off points as well and found similar results. This approach is intended primarily as a robustness test. Within Tables 24 and 25, the estimates differ somewhat, but a consistent picture emerges regardless of the measurement strategy used. Meso-structural sexism within marital dyads is associated with worse physical health among women, and with better physical and mental health among men. This supports a zero-sum gender conflict perspective.

**Table 24. Comparison of Meso-Sexism Results Using Alternative Specifications of Index Among Married Women**

	Conditions	Poor SRH	Poor Functioning
Original Index: H:W ratios, STD sum, wage ratio logged	0.090** (0.034)	0.031 (0.027)	1.336*** (0.284)
(1) H:W ratios, wages not logged	0.089* (0.037)	-0.001 (0.028)	0.781** (0.278)
(2) % of total attributable to H, STD sum	0.060 (0.038)	0.024 (0.026)	1.051*** (0.260)
(3) % of total attributable to H, sum	0.093** (0.033)	0.050* (0.023)	1.115*** (0.242)
(4) Dichotomized, sum	0.125** (0.041)	0.051 (0.030)	0.969** (0.314)
# of models with significant effects	4	1	5

Notes: H stands for husband, W for wife. Table shows X-standardized coefficient estimates (except for alternative 4, which is the raw coefficients) from random effects models among married women subsample. Robust standard errors are in parentheses. \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include macro- and micro-level sexism indices.

**Table 25. Comparison of Meso-Sexism Results Using Alternative Specifications of Index Among Married Men**

	Conditions	Poor SRH	Poor Functioning
Original Index: H:W ratios, STD sum, wage ratio logged	-0.052 (0.042)	-0.070** (0.024)	-0.709** (0.257)
(1) H:W ratios, wages not logged	-0.017 (0.042)	-0.023 (0.020)	0.096 (0.213)
(2) % of total attributable to H, STD sum	-0.072 (0.040)	-0.084** (0.026)	-0.939*** (0.237)
(3) % of total attributable to H, sum	-0.141*** (0.034)	-0.120*** (0.027)	-1.621*** (0.266)
(4) Dichotomized, sum	-0.030 (0.044)	-0.054 (0.029)	-0.867*** (0.221)
# of models with significant effects	1	3	4

Notes: H stands for husband, W for wife. Table shows X-standardized coefficient estimates (except for alternative 4, which is the raw coefficients) from random effects models among married men subsample. Robust standard errors are in parentheses. \* p<.05 \*\* p<.01 \*\*\* p<.001. All models include individual level controls (age, years of education, household income, race, divorce during observation period, health insurance, parental status, duration in state of residence) and state-level controls (poverty rate, Gini coefficient, racial composition, and southern region). Models do not include macro- and micro-level sexism indices.

One final consideration relevant for the creation of indices is the distinction between effect (or reflective) indicators and causal (or formative) indicators (Bollen 1989; Bollen and Bauldry 2011). Effect indicators are highly correlated and are assumed to change in response to the variation in some underlying latent construct. Causal indicators, on the other hand, may or may not be correlated but nonetheless contribute to the same latent construct. An example of this is the commonly used indices of stressful life events; a death in the family and getting married are unlikely to be correlated but may both be considered sources of stress (Bollen and Bauldry 2011). Similarly, a wife being younger than her husband may not be related to whether she earns less than him, but both can function as potential sources of inequality within the marriage. Thus, the meso-level structural sexism index is composed of causal indicators. Combining these indicators is important because inequality in any one of these indicators could be a source of power/resource inequality within the marriage, and we would expect that inequality in all three dimensions reflects an even greater power imbalance between spouses. At the macro-level, the structural sexism indicators are more highly correlated (Cronbach's  $\alpha = .64$ ) and the issue of whether the indicators are best conceptualized as causal vs. effect indicators is less clear-cut.

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## Biography

Patricia Homan was born on November 12, 1982 in Orlando, FL. She received a B.A. in Sociology with honors from Princeton University in 2005, and an M.A. in Sociology from Duke University in 2016. Her doctoral training and research at Duke University was supported by a grant from the National Institute on Aging (NIA grant # T32 AG000139) through the Duke Population Research Institute's T32 Pre-Doctoral Training Program in the Demography of Aging. She is the author/co-author of two peer-reviewed publications and one book chapter: "Political Gender Inequality and Infant Mortality in the United States, 1990-2012" (in *Social Science and Medicine*); "Being and Becoming Poor: How Cultural Schemas Shape Beliefs About Poverty" (with Lauren Valentino and Emi Weed in *Social Forces*); and "Health Coverage, Utilization, and Expenditures Among Latino Men" (with Russell Homan and Olveen Carrasquillo in *Health Issues in Latino Males: A Social and Structural Approach*). Her recent article entitled, "Rethinking the Role of Childhood SES in Adult Health: Integrating Existing Theories with Life Course Perspectives on the Disablement Process," (with Scott Lynch, currently under peer-review) received the Emerging Scholar Paper Award from the International Sociological Association's Research Committee on the Sociology of Aging.

Patricia will receive her Ph.D. from Duke University in May 2018. She will then continue her career as an Assistant Professor in the Department of Sociology and an Associate of the Pepper Institute on Aging and Public Policy at Florida State University.