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Political Gender Inequality and Infant Mortality in the United States, 1990–2012

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Abstract

Although gender inequality has been recognized as a crucial factor influencing population health in the developing world, research has not yet thoroughly documented the role it may play in shaping U.S. infant mortality rates (IMRs). This study uses administrative data with fixed-effects and random-effects models to (1) investigate the relationship between political gender inequality in state legislatures and state infant mortality rates in the United States from 1990 to 2012, and (2) project the population level costs associated with women's underrepresentation in 2012. Results indicate that higher percentages of women in state legislatures are associated with reduced IMRs, both between states and within-states over time. According to model predictions, if women were at parity with men in state legislatures, the expected number of infant deaths in the U.S. in 2012 would have been lower by approximately 14.6% (3,478 infant deaths). These findings underscore the importance of women's political representation for population health.

Keywords

United States; infant mortality; gender inequality; political representation

INTRODUCTION

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 26th 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.

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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 et al., 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 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.

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 et al., 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, 2016a). 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 (e.g. Lukachko et al., 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 et al., 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 et al., 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 et al., 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 et al., 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.

METHODS

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).

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, 2016b).

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, 2016b). 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,

2016c). 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 Table S1) 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 Table S1) adds a control for the political party affiliation of the sitting U.S. President (Rodriguez et al., 2013).

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^{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 $y_{it}^* = y_{it} - \bar{y}_i$, 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 Table S2). 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 Table S3). 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.

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 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.

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 Figure S1.)

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.

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.

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 Table S1). 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 Table S1). 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 Table S4). 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 Table S5). 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 28th out of 145 countries in overall gender equality in 2015, it ranked 72nd in women's political empowerment and 81st for the number of women in parliament (Hausmann et al., 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.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

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Research Highlights

- Finds strong link between political gender inequality and IMRs in the U.S.
- Projects the population level costs of women's underrepresentation in 2012
- Suggests women's political empowerment is important for U.S. population health

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

Figure 2.
Within-State Relationships between IMR and Women in Legislature, 1990–2012

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

1990				
	<u>Mean</u>	<u>SD</u>	<u>Min</u>	<u>Max</u>
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

Table 2

Zero-Order Correlations Between Key Variables

	IMR	Legislature % Women	Poverty	Nonwhite	Median Income	% HS Degree	Gini	South	Year
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	---	

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

* p<.05,

** p<.01,

*** p<.001, two-tailed tests.

Note: Standard errors are included in parentheses.