

SOCIAL SCIENCES

The evolution of infant mortality inequality in the United States, 1960–2016

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What is the relationship between infant mortality and poverty in the United States and how has it changed over time? We address this question by analyzing county-level data between 1960 and 2016. Our estimates suggest that level differences in mortality rates between the poorest and least poor counties decreased meaningfully between 1960 and 2000. Nearly three-quarters of the decrease occurred between 1960 and 1980, coincident with the introduction of antipoverty programs and improvements in medical care for infants. We estimate that declining inequality accounts for 18% of the national reduction in infant mortality between 1960 and 2000. However, we also find that level differences between the poorest and least poor counties remained constant between 2000 and 2016, suggesting an important role for policies that improve the health of infants in poor areas.

INTRODUCTION

In contrast to growing disparities in life expectancy between the rich and poor (1–9) and increasing economic inequality (10–12), recent studies find decreasing inequality in mortality and health outcomes among younger age groups in the United States (13, 14). While there is a general understanding that poverty is associated with worse childhood health outcomes, including mortality (15), less is known about how such inequality evolved over a longer period.

Here, we aim to understand the evolution of inequality in infant mortality between 1960 and 2016. We ask: How long have infant mortality rates been converging between the rich and poor? How does the reduction in mortality inequality relate to the overall decline in national infant mortality rates?

We address these questions by analyzing the association between infant mortality and poverty in the United States between 1960 and 2016. Our analysis sample uses county-level data on mortality rates from the U.S. Centers for Disease Control and Prevention (CDC) and poverty data from the U.S. Census Bureau. We construct 1-year rates (per 1000 live births) for infant mortality (death under 1 year), neonatal mortality (death under 28 days), and postneonatal mortality (death between 28 and 364 days) at the county-year level, and define poverty ventiles as 20 equally sized (by population) aggregations of counties in each year. We then conduct the following analyses on infant, neonatal, and postneonatal mortality rates: (i) quantify the association between mortality rates and poverty rank using the mortality-poverty gradient and (ii) decompose national mortality rate declines to approximate the share accounted for by reduced mortality inequality. Following Currie and Schwandt (14), our methodology focuses on the evolution of level differences in infant mortality rates over the poverty distribution (although we also present results on percentage differences in Materials and Methods).

RESULTS AND DISCUSSION

Over the last 50 years, mean poverty and infant mortality rates fell considerably in the United States. The average poverty rate decreased

from 22.1% in 1960 to 14.7% in 2016, while mean infant mortality fell from 25.8 in 1960 to 5.8 in 2016. Level differences in poverty and infant mortality also narrowed considerably between the most poor and least poor counties. As shown in Table 1, the difference in average poverty rates between the poorest 5% of counties (top poverty ventile) and the least poor 5% (bottom poverty ventile) declined from 55.1 percentage points in 1960 to 22.0 percentage points in 2016 ($P = 0.000$). Over this period, there was a relatively larger reduction in the poverty rate among the poorest counties, which may be due to targeted federal spending (16) enacted in the 1960s and 1970s as part of the War on Poverty. Level differences in infant mortality rates also narrowed between the most poor and least poor counties, declining from 17.8 in 1960 to 3.6 in 2016 ($P = 0.000$). The differences in both neonatal and postneonatal mortality rates between the most poor and least poor counties also narrowed meaningfully over this period.

Figure 1 explores the relationship between poverty rank and infant mortality rates for 1960, 1970, 1980, 1990, 2000, 2010, and 2016. Figure 1A shows results for infant mortality, Fig. 1B shows results for neonatal mortality, and Fig. 1C shows results for postneonatal mortality. The lines in each figure show fitted values from Eq. 1 that relate the given mortality rate to the poverty ventile based on population-weighted, county-level regressions estimated for each year. The data points shown in the figure plot the average mortality rate in each poverty ventile in 1960 and 2016 (shown only for these years to improve readability of the figure). Figure 1 suggests two key patterns. First, poverty rank is positively correlated with mortality rates in all years, as evident from the upward slope of the fitted lines. Second, the overall national declines in mortality rates between 1960 and 2016 mask substantial heterogeneity across the poverty distribution.

To illustrate this second key pattern, Fig. 2 plots mortality-poverty gradients, highlighting our first main conclusion: Inequality in infant mortality, as measured by the association between infant mortality rates in levels and poverty rankings, declined significantly between 1960 and 2000 but was statistically unchanged from 2000 to 2016. The mortality-poverty gradients quantify the association between a given mortality rate measure and a one-unit increase in the poverty ventile distribution. For infant mortality (Fig. 2A), the gradient decreased markedly from 0.581 in 1960 to 0.141 in 2000 ($P = 0.000$). As measured by the mortality-poverty gradient, the level of mortality inequality in 2000 was roughly 25% of the 1960 level. Despite a continuing decline in the national infant mortality rate between

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Table 1. Mean poverty rates and mortality rates, by poverty ventile, 1960 and 2016. Poverty ventiles (equally sized groups representing 5% of the population) are defined in each year based on county-level data using population estimates from the CDC and poverty rates from the U.S. Census Bureau. Means are calculated as the population-weighted average across all counties in each poverty ventile and are reported for 1960 and 2016. Columns 2 and 3 report the average poverty rate. The remainder of the columns report 1-year mortality rates (expressed as deaths per 1000 births). Columns 4 and 5 show the infant mortality rate (deaths before age 1), columns 6 and 7 report neonatal mortality rate (deaths from 0 to 28 days), and columns 8 and 9 report postneonatal mortality rate (deaths from 29 to 364 days). Data for mortality rates are from the CDC, and data for poverty rates are from the U.S. Census Bureau.

Poverty ventile	Poverty rate		Infant mortality rate		Neonatal mortality rate		Postneonatal mortality rate	
	1960	2016	1960	2016	1960	2016	1960	2016
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1	6.7	5.8	20.4	4.0	16.0	2.8	4.3	1.2
2	8.8	7.4	21.8	4.2	16.8	3.1	4.9	1.1
3	10.3	8.6	21.9	4.8	16.6	3.3	5.3	1.5
4	11.9	9.6	23.8	5.0	18.4	3.5	5.4	1.6
5	12.5	10.5	24.2	5.0	17.7	3.3	6.5	1.7
6	13.0	11.3	23.3	5.4	17.5	3.9	5.8	1.5
7	13.7	12.2	23.9	5.6	18.1	3.8	5.8	1.9
8	14.6	12.9	24.0	5.2	17.9	3.3	6.1	1.9
9	15.8	13.7	24.0	5.1	18.2	3.3	5.7	1.8
10	16.7	14.3	25.0	6.1	18.7	4.0	6.3	2.2
11	17.6	15.2	24.4	6.5	18.1	4.1	6.3	2.4
12	19.2	15.8	26.1	6.6	19.6	4.5	6.4	2.1
13	21.3	16.2	24.5	6.2	17.9	4.3	6.6	1.9
14	23.6	16.6	27.3	6.9	20.1	4.5	7.2	2.4
15	26.0	16.9	26.6	5.7	19.3	3.9	7.3	1.8
16	29.5	17.7	26.6	6.1	19.0	4.0	7.6	2.0
17	32.9	18.8	28.1	7.3	19.6	4.7	8.5	2.6
18	38.1	20.1	29.5	6.9	19.6	4.3	9.9	2.5
19	47.3	22.4	32.1	7.0	20.5	4.4	11.6	2.6
20	61.8	27.8	38.2	7.6	22.3	4.7	16.0	2.9

2000 and 2016 (from 6.1 to 5.8), the gradient increased over this period from 0.141 to 0.153, although not significantly ($P = 0.737$). Intuitively, while the overall infant mortality rate continued to decline between 2000 and 2016, the differential mortality rate reductions among the poorest counties, relative to the least poor counties, did not. We interpret the gradient as unchanged in these later years given the small magnitude of the increase and the lack of statistical significance. Table 2 reports gradient estimates and SEs.

Both neonatal and postneonatal mortality gradients follow a similar pattern of large declines between 1960 and the early 2000s, with statistically insignificant increases in more recent years. For neonatal mortality (Fig. 2B), the gradient dropped from 0.218 in 1960 to 0.066 in 2010 ($P = 0.000$) but then increased from 0.066 to 0.083 between 2010 and 2016 ($P = 0.180$). For postneonatal mortality (Fig. 2C), the gradient dropped from 0.363 in 1960 to 0.074 in 2000 ($P = 0.002$) and then increased to 0.079 in 2010 ($P = 0.704$) before declining again to 0.070 in 2016 ($P = 0.380$).

Figure 2 also suggests that inequality reductions occurred largely between 1960 and 1980 and that declines in postneonatal infant mortality inequality explain most of the overall inequality reduction in infant mortality. In Fig. 2A, 73% of the decline between 1960 and

2000 occurred between 1960 and 1980. Decomposing the infant mortality gradient (Fig. 2A) into the neonatal (Fig. 2B) and postneonatal (Fig. 2C) gradients shows that 25% of the 1960–1980 decline in the infant mortality gradient is explained by neonatal mortality, whereas 75% is explained by postneonatal mortality.

Our findings are consistent with prior work analyzing infant mortality inequality over shorter time periods and with studies analyzing specific social programs. We show that the weakening association between poverty and the level of infant mortality reported by Currie and Schwandt (14) began before 1990 and that declining inequality is found for both neonatal and postneonatal mortality rates. The pattern of sharp reductions in mortality inequality that we document, starting in the 1960s and tapering off in later years, is also broadly consistent with Singh and Kogan (17). Compared to these studies (14, 17), our analysis covers a wider range of years and is notable for documenting the flattening of the gradient between 2000 and 2016. The large mortality rate reductions we find among the poorest counties between 1960 and 1980 are consistent with work that reports positive effects on infant health from federal antipoverty programs that began during those decades, including food stamps (18), Medicaid (19), and community health centers (20), as well as work that documents improvements in medical

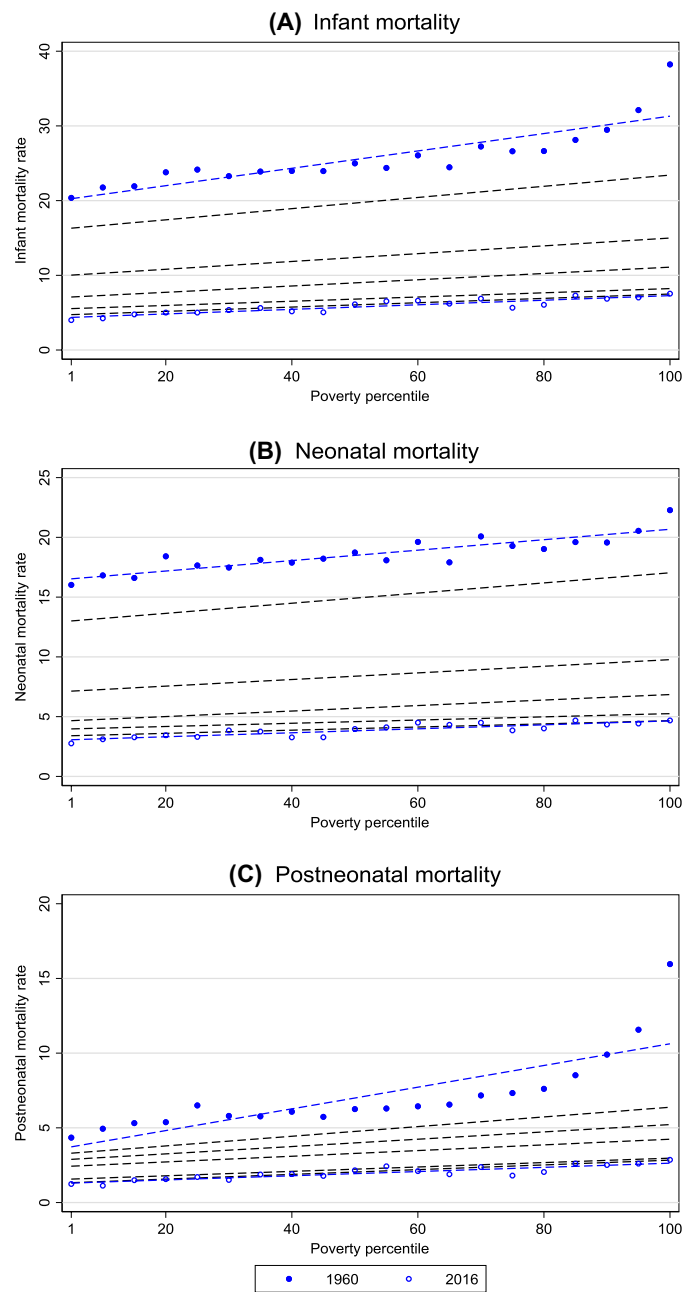


Fig. 1. Mortality rates and fitted values. (A) shows results for infant mortality rates, (B) shows results for neonatal mortality rates, and (C) shows results for postneonatal mortality rates (mortality rates defined in the notes to Table 1). The lines in each figure show fitted values estimated from population weighted county-level regressions each year using Eq. 1 and are shown for 1960, 1970, 1980, 1990, 2000, 2010, and 2016. Table 2 shows parameter estimates for each regression. Data points in the figures show mean mortality rates for each poverty ventile for 1960 and 2016 only to improve readability of the figures. Data on mortality rates are from the CDC, and data on poverty rates are from the U.S. Census Bureau.

care for infants during this time period (21, 22). Last, the key role of postneonatal mortality in explaining the infant mortality gradient reduction is consistent with work showing that socioeconomic disparities emerge especially strongly during the postneonatal period (23).

We explore heterogeneity by gender in the evolution of infant mortality inequality in Fig. 3, which suggests that the raw gap in infant mortality rates between males and females narrowed between 1970 and 2016 but that changes in the level of infant mortality inequality were comparable across genders. Our data on gender-specific mortality rates start in 1970. Nationally, the infant mortality gap across genders decreased from 4.8 in 1970 to 1.0 in 2016 ($P = 0.000$), with continual declines over the sample period. Figure 3A shows this narrowing gender gap over time, plotting fitted values from Eq. 1 and mean mortality rates by poverty ventiles in select years. This result is consistent with work showing that excess male infant mortality generally peaked in the 1970s across 15 advanced economies (24) and with results showing a narrowing of the infant mortality gender gap in Massachusetts in the early 1990s (25). Mirroring the national pattern in Fig. 2, both male and female infant mortality gradients declined from 1970 through 2000 but then remained flat thereafter (Fig. 3B). As shown in Table 2, the male (female) gradient fell significantly from 0.406 (0.338) in 1970 to 0.150 (0.132) in 2000 and then increased insignificantly to 0.160 (0.145) in 2016. Although our gradient estimates imply a relatively larger reduction in the level of infant mortality inequality for males, the differences in the gradients across genders are not statistically different in any year, and we therefore cannot rule out equal decreases.

To highlight the association between changes in the mortality-poverty gradient and national mortality rates, we construct a set of counterfactual mortality rates. We define these counterfactual measures as the national mortality rates that would have occurred in later years if the mortality-poverty gradient from a given base year were held fixed and each poverty ventile experienced the same change in mortality as the least poor ventile. We set the base year to 1960 for aggregate results (1970 for gender-specific results). The counterfactual rates rest on two strong assumptions: (i) that mortality inequality is unchanged from the base year level and (ii) that each poverty ventile has the same trend as the least poor ventile over time. Instead of representing a likely scenario, we view the counterfactual rates as helpful benchmarks. The counterfactual poses the question: How would mortality rates have evolved nationally if each county in the United States experienced the same change as the least poor ventile, thus holding fixed the level of mortality inequality? The difference between the actual and counterfactual rates reflects the additional mortality rate reductions among higher poverty ventiles, compared to the least poor ventile, thus highlighting the role of mortality inequality reductions.

Figure 4 shows the actual (solid lines) and counterfactual (dashed lines) infant mortality rates, yielding our second main conclusion: The reduction in mortality inequality accounts for a meaningful share of the overall decline in the national infant mortality rate during our period. Compared to the actual reduction in infant mortality, the reduction in the counterfactual infant mortality is 18% smaller between 1960 and 2016 (18% smaller between 1960 and 2000). Correspondingly, by 2016 (2000), the national infant mortality rate would have been 1.6 (1.5) times larger in the counterfactual compared to the actual. The pattern is similar for neonatal (postneonatal) in Fig. 4B (Fig. 4C), with reductions that would have been 10% (40%) smaller between 1960 and 2016. This pattern is also similar by gender, with smaller counterfactual reductions of 18% (14%) for female (male) infant mortality rates between 1970 and 2016 (table S1).

Table 2. Mortality-poverty gradient estimates. (A) shows results for infant mortality, (B) shows results for neonatal mortality, and (C) shows results for postneonatal mortality. Each coefficient and SE (shown in brackets) is from a separate regression and represents the effect on the given mortality rate of increasing one unit on the poverty ventile distribution, estimated using Eq. 1. Gender-specific mortality rates are available in 1970 and later years only. SEs are clustered at the poverty ventile level. Data on mortality rates are from the CDC, and data on poverty rates are from the U.S. Census Bureau.

Sample	1960	1970	1980	1990	2000	2010	2016
(A) Infant mortality							
Entire sample	0.581 [0.111]	0.374 [0.052]	0.261 [0.025]	0.210 [0.026]	0.141 [0.031]	0.145 [0.014]	0.153 [0.014]
<i>N</i>	3096	3104	3104	3128	3134	3136	3137
Females		0.338 [0.053]	0.248 [0.025]	0.187 [0.025]	0.132 [0.030]	0.133 [0.018]	0.145 [0.017]
<i>N</i>		3046	3103	3128	3132	3136	3134
Males		0.406 [0.058]	0.276 [0.030]	0.234 [0.030]	0.150 [0.037]	0.156 [0.015]	0.160 [0.016]
<i>N</i>		3043	3104	3127	3134	3135	3135
(B) Neonatal mortality							
Entire sample	0.218 [0.031]	0.212 [0.020]	0.139 [0.016]	0.115 [0.022]	0.067 [0.020]	0.066 [0.008]	0.083 [0.008]
<i>N</i>	3096	3104	3104	3128	3134	3136	3137
Females		0.181 [0.024]	0.130 [0.016]	0.098 [0.019]	0.056 [0.019]	0.064 [0.012]	0.074 [0.012]
<i>N</i>		3046	3103	3128	3132	3136	3134
Males		0.238 [0.023]	0.148 [0.020]	0.132 [0.027]	0.078 [0.024]	0.067 [0.009]	0.091 [0.009]
<i>N</i>		3043	3104	3127	3134	3135	3135
(C) Postneonatal mortality							
Entire sample	0.363 [0.086]	0.162 [0.038]	0.123 [0.012]	0.095 [0.008]	0.074 [0.013]	0.079 [0.007]	0.070 [0.008]
<i>N</i>	3096	3104	3104	3128	3134	3136	3137
Females		0.158 [0.035]	0.118 [0.013]	0.089 [0.009]	0.076 [0.012]	0.070 [0.009]	0.072 [0.009]
<i>N</i>		3046	3103	3128	3132	3136	3134
Males		0.168 [0.044]	0.127 [0.012]	0.101 [0.009]	0.072 [0.015]	0.088 [0.008]	0.069 [0.009]
<i>N</i>		3043	3104	3127	3134	3135	3135

CONCLUDING REMARKS

Alleviating inequality in health generally, and in mortality specifically, remains a major policy goal (26). Yet, the extent to which inequality in infant mortality has changed over the past 50 years is not well documented. We analyze data from the CDC and U.S. Census Bureau to address these issues and to establish new facts about the association between infant mortality and poverty over time.

Our findings suggest both positive developments and concerning patterns. In terms of positive developments, we show that infant mortality rates have become more equal across the poorest and least poor counties in the United States between 1960 and 2000, with our estimate of mortality inequality declining 76%. We estimate that this reduction in inequality accounts for nearly 20% of the overall

national reduction between 1960 and 2000. In addition, we show that the gender gap in infant mortality rates narrowed substantially between 1970 and 2016. Yet, we also document that mortality inequality has remained roughly constant between 2000 and 2016 and that the level of inequality that remained in 2016 was roughly one-quarter of the 1960 level. The trend in recent years, together with the persisting level of inequality, suggests an important role for policies that improve the health of infants in poor areas.

Our study has several limitations. First, the analysis of the mortality-poverty gradients uses data aggregated to the county level. As a result, the findings do not represent comparisons of poor versus rich individuals; rather, they represent the association between mortality and residing in an area with a higher poverty rate compared to an area with a lower poverty rate. Second, our

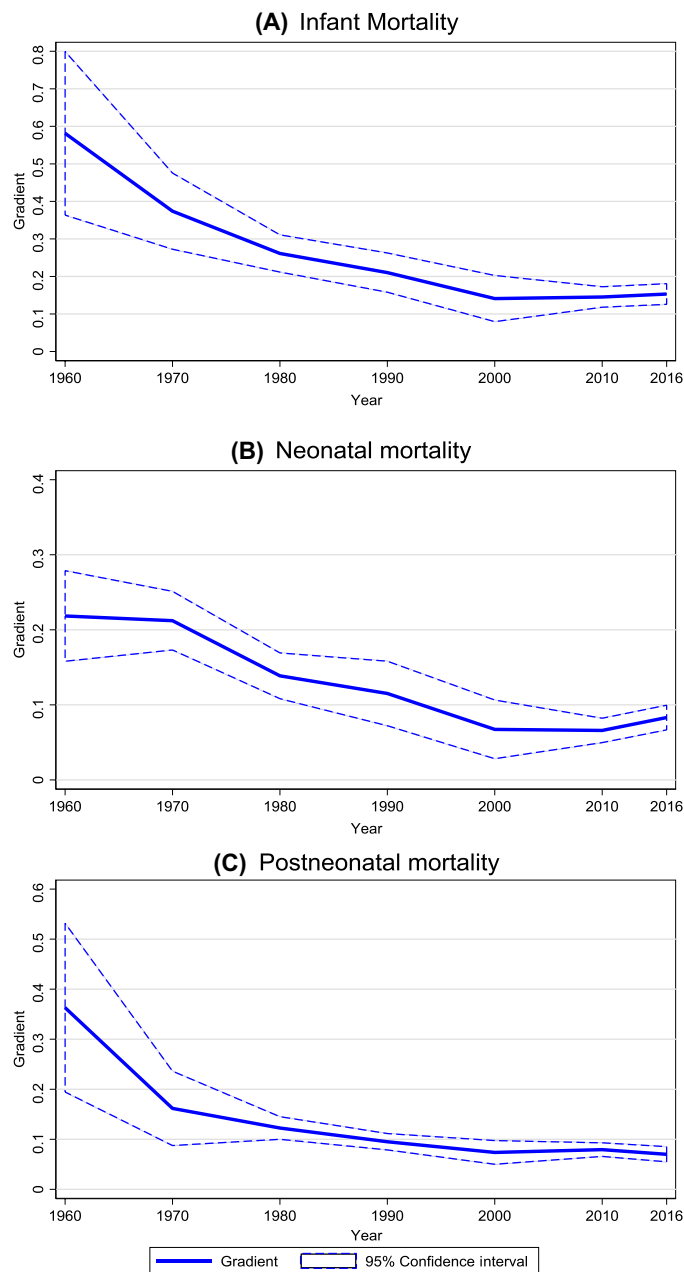


Fig. 2. Mortality-poverty gradient estimates. (A) shows results for infant mortality, (B) shows results for neonatal mortality, and (C) shows results for postneonatal mortality. The lines in each figure show parameter estimates of the effect of increasing one unit in the poverty ventile distribution on the given mortality rate, estimated in each year using Eq. 1. The shaded areas show the 95% confidence intervals based on SEs clustered at the poverty ventile level. Table 2 reports parameter estimates and SEs. Data on mortality rates are from the CDC, and data on poverty rates are from the U.S. Census Bureau.

specification uses only variation in poverty across counties. It therefore does not account for inequality within a county, which could be an interesting area for future work. Third, our measure of inequality uses only the poverty rate, which overlooks other potentially important features of a county’s income distribution, such as median income or income inequality. We selected this measure because it allows us to rank all counties along a single

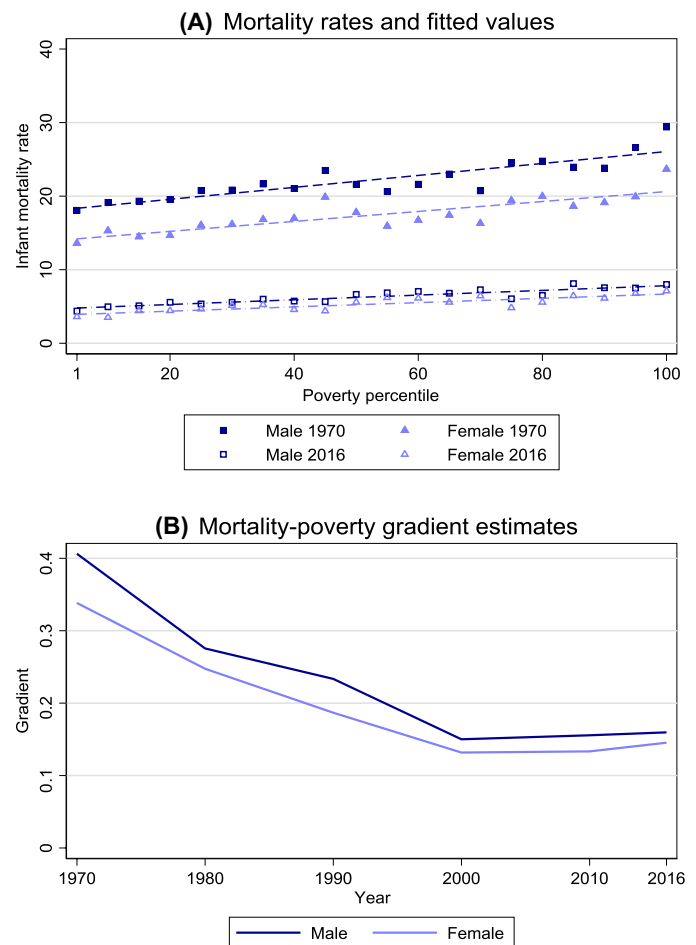


Fig. 3. Mortality rates, fitted values, and gradients, by gender. Both panels show results for infant mortality. Because of data limitations, the first available year of data for these figures is 1970. The lines in (A) and the gradients in (B) are based on estimates from weighted county-level regressions each year using Eq. 1. Table 2 reports parameter estimates for each regression. In (A), data points show mean mortality rates for each poverty ventile for 1970 and 2016 only to improve readability of the figures. In (B), we omit the confidence intervals to increase readability of the figure as the confidence intervals overlap in each year. Data on mortality rates are from the CDCs, and data on poverty rates are from the U.S. Census Bureau.

dimension in a transparent way. Fourth, our estimates of the mortality-poverty gradients cannot be interpreted as the causal effect of residing in poor versus rich areas. A county’s poverty rate is likely correlated with a variety of factors that directly affect mortality in the first year of life. Because of these confounding factors, the causal effect of poverty on mortality is likely to be different from the estimates in this study.

MATERIALS AND METHODS

Data

To measure infant mortality, we use data from the CDC, and to measure poverty, we use data from the U.S. Census Bureau. We construct 1-year mortality rates (per 1000 live births) for infant mortality (death under 1 year), neonatal mortality (death under

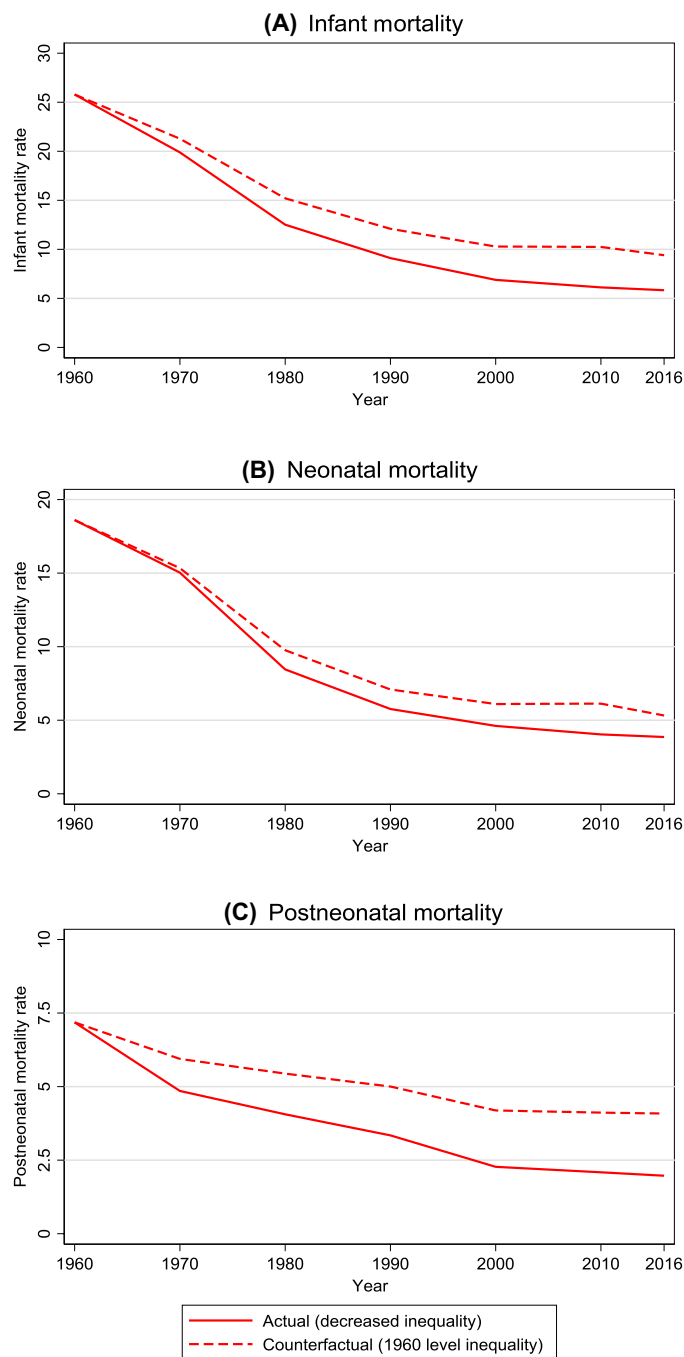


Fig. 4. Counterfactual and actual mortality rates, by year. (A) shows results for infant mortality rates, (B) shows results for neonatal mortality rates, and (C) shows results for postneonatal mortality rates (mortality rates defined in the notes to Table 1). The solid lines show actual mortality rates. The dashed lines show counterfactual mortality rates, defined as the mortality rate that would exist if (i) the mortality gradient remained at the 1960 value reported in Table 2 and (ii) each poverty ventile experienced a reduction in mortality rates over time equal to the actual reduction experienced by the least poor ventile. Table S1 reports the counterfactual estimates in each year. Data on mortality rates are from the CDC, and data on poverty rates are from the U.S. Census Bureau.

28 days), and postneonatal mortality (death between 28 and 364 days) at the county-year level, corresponding to the years 1960, 1970, 1980, 1990, 2000, 2010, and 2016. Data on poverty rates were matched to each county-year observation using county of residence. Poverty ventiles, 20 equally sized (by population) groups of counties, were defined in each year based on rankings of the poverty rate. Although it need not be the case, each poverty ventile also represents about 5% of births each year. In table S2, we show that our results are robust to alternate samples of counties, including restricting to counties that do not undergo boundary changes over time and to the set of counties available in all years. In table S3, we show that the gender-specific results are similar when restricting to the set of counties that have nonmissing mortality rates for both genders. In table S4, we show results that hold fixed the poverty ranking of counties from 1960, which are similar to the results from our baseline approach that ranks counties in each year.

Our analysis sample consists of 21,839 county-year level observations across the years 1960, 1970, 1980, 1990, 2000, 2010, and 2016 (18,743 county-year observations for 1970 and later years). Our data include 31,606,403 births and 365,171 infant deaths (254,321 neonatal deaths and 110,850 postneonatal deaths) between 1960 and 2016. For gender-specific outcomes, our sample includes 14,008,280 male births, 13,344,195 female births, 144,195 male infant deaths (98,901 male neonatal deaths and 45,294 male postneonatal deaths), and 110,834 female infant deaths (76,137 female neonatal deaths and 34,697 female postneonatal deaths) for 1970 and later years.

We report results for race-specific mortality in table S5, although we urge caution when interpreting these findings because of two important data limitations. First, the method of assigning race at birth changes over time. Before 1989, race assignment at birth used information from both parents; thereafter, it was based on the race of the mother, as stated on the birth certificate. This reclassification results in more births classified as white and fewer births classified as nonwhite, causing white infant mortality rates to drop and nonwhite infant mortality rates to increase (27, 28). A second concern is error in race assignment at death, which may be reported by family members, or by the funeral director based on observation, and thus has differential error by race (29, 30).

Baseline specification

To quantify the relationship between poverty rank and infant mortality, we use ordinary least squares regression analysis of Eq. 1, following Currie and Schwandt (14)

$$mortality_{ct} = \alpha + \beta * poverty_{ct} + \varepsilon_{ct} \quad (1)$$

where c denotes the county and t denotes the year. Poverty is defined as the poverty ventile (1 to 20). Dependent variables include the following mortality rates: infant, neonatal, and postneonatal. The ε_{ct} term captures model error and α is the intercept. We weight each county observation by population and estimate the equation separately for each year, which is equivalent to a model that includes year indicator variables and interactions of year and poverty rank. We cluster SEs at the poverty ventile level to allow for arbitrary correlation in the error structure within poverty groups.

The coefficient β , referred to as the mortality-poverty gradient, reflects the difference in the level of mortality rates per unit increase in the poverty ventile (or per 5 percentile points in the poverty

ranking). A zero value for this gradient would indicate equality in mortality rates across the poverty distribution. Positive values would indicate inequality, reflecting lower mortality rates in the least poor counties compared to the poorest counties. These results can be interpreted as quantifying absolute inequality, as they relate the level of mortality rates and the poverty rank. Below, we present results for relative inequality measures that relate percent changes in mortality rates and poverty rank. We then discuss our decision to focus on absolute mortality inequality measures in the main text.

Additional specifications

We also estimate the mortality-poverty rate gradient using Eq. 1a, which quantifies absolute inequality between the mortality rate and the poverty rate

$$mortality_{ct} = \alpha + \gamma * poverty_{rate}_{ct} + \varepsilon_{ct} \quad (1a)$$

The findings are roughly similar when using the poverty rate rather than the poverty rank. The results using the poverty rate are shown in figs. S1 and S2 and table S6. We use the poverty rank in our baseline method due to well-known concerns about comparing poverty rates over time (31) and because aggregate measures of income at the county level may not capture the dynamics of income inequality driven by rapid growth at the top of the distribution (10, 11).

To estimate the shift in the poverty rate distribution over time, we estimate Eq. 2, which relates the poverty rate to the poverty rank and present these results in panel D of table S6.

$$poverty_{rate}_{ct} = \alpha + \lambda * poverty_{ct} + \varepsilon_{ct} \quad (2)$$

Last, to quantify inequality in percentage terms (i.e., relative inequality), we report the results for transformations of the mortality outcomes using Eq. 3

$$f(mortality_{ct}) = \alpha + \phi * poverty_{ct} + \varepsilon_{ct} \quad (3)$$

We use several functions $f(-)$ of mortality rates, including the natural log transformation that relates the percent change in the mortality rate to the poverty rank. However, this method excludes counties with zero values of mortality rates because the natural log is not defined in these cases. There are roughly 100 such counties (0.2% of counties population weighted) in 1960, with the number generally increasing each year to 822 (2.7% of counties population weighted) in 2016. To address the issue of omitting counties with a zero mortality rate, we include additional transformations frequently used when encountering zero values (32). First, we use the natural log transformation and include counties with zero mortality rate by assigning them a zero value for the transformation. Second, we use the inverse hyperbolic sine that also facilitates a percent change interpretation and is defined for zero values. Third, we analyze regressions at the poverty ventile level that use the natural log of the mean infant mortality rate in each ventile.

The results for these transformations are shown in table S7. The point estimates generally follow an upward trend. However, the sizes of the SEs limit the extent to which we can make statistically significant comparisons across years. We interpret these results as suggesting that relative inequality in mortality was constant or perhaps increasing over our time period. The different interpretations on the evolution of infant mortality inequality when we analyze mortality in logs rather than levels stem from changes in the mortality rate

over time. Conceptually, the natural log specification scales the parameter estimates from Eq. 1 by the underlying mean mortality rate in each year. Because both the level differences and the mortality rate decline roughly proportionally over time, the log differences follow a constant or slightly increasing trend. Overall, our results on relative mortality inequality raise concerns about infant health disparities over our sample period and serve to strengthen our conclusion that improving the health of all infants, but especially those in poor areas, should be a key priority for policies in the coming years.

We use the mortality rate in levels as our baseline measure because it focuses on absolute rather than relative changes. The absolute change in mortality rates reflects the number of deaths, independent of the pre-existing infant mortality level. To highlight our rationale for focusing on mortality rate levels, consider the following example modified from Currie and Schwandt (14). In the base period, the mortality rate is 2 in the least poor county and 20 in the poorest county, while in the subsequent period, the mortality rate decreases by 1 in the least poor county and by 9 in the poorest. On the basis of relative changes, the decrease is larger in the least poor county (50% versus 45%), but in absolute terms, the decrease is much larger in the poorest county (9 versus 1). We thus focus on absolute changes in the infant mortality rates as this method weights deaths in both the least poor and the most poor counties the same.

All models were estimated using STATA 16.0 MP with the regress command, analytical weights set to county-level population, and SEs clustered at the poverty ventile level. Throughout the paper, we use two-tailed tests of statistical significance at the 0.05% level.

SUPPLEMENTARY MATERIALS

Supplementary material for this article is available at <http://advances.sciencemag.org/cgi/content/full/6/29/eaba5908/DC1>

REFERENCES AND NOTES

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