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# Other Articles

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## Is Exposure to Income Inequality a Public Health Concern? Lagged Effects of Income Inequality on Individual and Population Health

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**Objective.** To examine the health consequences of exposure to income inequality.

**Data Sources.** Secondary analysis employing data from several publicly available sources. Measures of individual health status and other individual characteristics are obtained from the March Current Population Survey (CPS). State-level income inequality is measured by the Gini coefficient based on family income, as reported by the U.S. Census Bureau and Al-Samarrie and Miller (1967). State-level mortality rates are from the *Vital Statistics of the United States*; other state-level characteristics are from U.S. census data as reported in the *Statistical Abstract of the United States*.

**Study Design.** We examine the effects of state-level income inequality lagged from 5 to 29 years on individual health by estimating probit models of poor/fair health status for samples of adults aged 25–74 in the 1995 through 1999 March CPS. We control for several individual characteristics, including educational attainment and household income, as well as regional fixed effects. We use multivariate regression to estimate the effects of income inequality lagged 10 and 20 years on state-level mortality rates for 1990, 1980, 1970, and 1960.

**Principal Findings.** Lagged income inequality is not significantly associated with individual health status after controlling for regional fixed effects. Lagged income inequality is not associated with all cause mortality, but associated with reduced mortality from cardiovascular disease and malignant neoplasms, after controlling for state fixed-effects.

**Conclusions.** In contrast to previous studies that fail to control for regional variations in health outcomes, we find little support for the contention that exposure to income inequality is detrimental to either individual or population health.

**Key Words.** Income inequality, social determinants of health, health status, mortality

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There is a sharp division among health policy researchers regarding the extent to which income inequality is a public health problem: some advocate the health benefits of egalitarian social policies (e.g., Wilkinson 1996, and Daniels, Kennedy, and Kawachi 2000), while others caution that there is little credible

evidence that inequality is a determinant of public health (e.g., Milyo and Mellor 1999; Deaton 2001a; and Mackenbach 2002).<sup>1</sup> The main point of contention is how to weigh the relative merits of empirical studies that report findings consistent with the claim that inequality has important and deleterious consequences for individual and population health (e.g., Lochner et al. 2001, and the studies collected in Kawachi, Kennedy; and Wilkinson 2000) against several recent empirical studies that find no such connection between inequality and health (e.g., Meara 1999; Deaton 2001b; Deaton and Lubotsky 2002; Deaton and Paxson 2001; Mellor and Milyo 2001 and 2002; Miller 2001; Gravelle, Wildman, and Sutton 2002; and Sturm and Gresenz 2002). In our opinion, two important methodological improvements in these latter studies are the inclusion of more detailed information on individual characteristics as control variables and the use of procedures that control for unobservable fixed-state or regional effects. Nevertheless, research studies on both sides of this literature share a common shortcoming.

With only two exceptions, every study of which we are aware investigates the contemporaneous association between income inequality and health outcomes, or the contemporaneous association between changes in inequality and changes in health outcomes. This is surprising, since the pathways by which inequality is conjectured to affect health clearly involve some persistent exposure, or seem to suggest a chain of events that might span years. For example, Wilkinson (1996) has argued that income disparities create psychosocial stress that eventually leads to heart disease or other stress-related maladies, while Kawachi et al. (1997) have speculated that inequality erodes social capital, which in turn eventually creates a political climate that is less supportive of policies that would improve public health. Consequently, explorations of the contemporaneous association between income inequality and health may fail to show evidence of a relationship that requires sufficient time to observe, and the question of whether prolonged exposure to income inequality is a determinant of either individual or population health remains unanswered by most of the extant quantitative research.

Blakely et al. (2000 and 2002) are the only attempts to date to address this shortcoming. The former study examines whether lagged and contemporaneous levels of state-level inequality (measured by the Gini coefficient) are

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associated with self-reported health status in a pooled cross-section of respondents to the Current Population Survey (CPS) in 1995 and 1997. The latter is a follow-on study that examines the association between lagged metropolitan area inequality and self-reported health status using CPS data for 1996 and 1998. Blakely et al. (2002) find that the association between metropolitan area inequality and health status is much weaker than for state-level income inequality; this is consistent with what has been found for contemporaneous associations between inequality and health status in the CPS (Mellor and Milyo 2002). In addition, the apparent importance of state inequality supports the causal pathway articulated by Kawachi et al. (1997): inequality erodes social capital, which in turn hinders political support for policies that improve public health.

Blakely et al. (2000) argue that lagged state inequality is more strongly associated with health status than contemporaneous inequality. For individuals older than 44 years of age, they find that the association between health status and the state-level Gini coefficient is strongest when the inequality measure is lagged 15 years, although they cannot reject the null hypothesis that the association is identical to that found with either shorter lags or the contemporaneous Gini, nor do they find any significant association for individuals younger than 45 years of age. Nevertheless, the authors conclude that these results are suggestive that exposure to inequality 15 years prior is an important determinant of current health status.

We agree that these findings are intriguing, but unfortunately the model on which they are based does not control for many of the same factors omitted in the earlier literature. When included in more recent studies, these factors produce results that do not support the claim that inequality adversely affects health. For example, Blakely et al. (2000) do not control for certain individual covariates that are available in the CPS (e.g., education, marital status, and health insurance coverage). More importantly, the authors do not control for regional variations in health status, despite the fact that even textbook accounts of population health in the United States note the importance of regional differences in a variety of inputs in the health production function, such as lifestyle (e.g., Phelps 1997). This oversight is problematic because of the fact that income inequality is also known to be related to factors that vary regionally across the United States. For example, the size and age-composition of households, the prevalence of households with one adult or one working adult, immigration, and the number of jobs in the manufacturing sector have all been documented to be determinants of regional variations in income inequality (Husted 1991; Levernier, Rickman, and Partridge 1995; Partridge,

Partridge, and Rickman 1998; Bernard and Jensen 2000). The failure to control for these regional effects may generate biased estimates of the effect of state-level income inequality on either individual or population health (Mellor and Milyo 2002).

In this article, we test whether the adverse consequences of lagged inequality are robust to the inclusion of other individual characteristics, and to the control of regional factors. In addition, we examine the strength of this association in multiple years of the CPS, using lags ranging from 5 to 29 years. In the second part of this article we use state-level data on all-cause mortality and cause-specific death rates to examine the lagged effect of income inequality on other health outcomes. Our results indicate that once we account for regional differences, there is insufficient evidence to support the claim that exposure to income disparities poses a risk to individual or population health.

## DATA AND METHODS

We conduct two separate analyses, one using individual-level data and one using state-level data. The data and methods in the individual analysis follow that in Mellor and Milyo (2002), while the data and methods in the state-level analysis are similar to that in Mellor and Milyo (2001). The important distinction of this study compared to our other work is that the explanatory variables of interest are various lagged values of the state-level Gini coefficient. These measures are calculated using observations for family income from decennial census data, and were obtained for this study from the U.S. Census Bureau web site ([www.census.gov/hhes/income/histinc/state/state4.html](http://www.census.gov/hhes/income/histinc/state/state4.html)) and Al-Samarrie and Miller (1967).

### *Individual-Level Analysis*

All individual-level data are from the March CPS, years 1995 through 1999.<sup>2</sup> Each year's sample includes white and black civilians not living in group quarters, between the ages of 25 and 74, and without missing information on key covariates. This leaves us with an average of almost 62,000 observations per year. The dependent variable in this analysis is a dichotomous measure of health status that takes the value of 1 if respondents rate health status as fair or poor, and 0 otherwise.<sup>3</sup> Roughly 14 percent of all respondents in these annual samples report fair or poor health status (see appendix Table A1 for these and other descriptive statistics).

To test the hypothesis that income inequality has long-term effects on health status, we estimate the marginal effects of state-level inequality on the probability that individual health status is fair or poor. In each model, we include control variables for age, age squared, race, Hispanic ethnicity, highest level of education completed, sex, residence in a metropolitan area or central city area, marital status, and health insurance coverage status (either public or private). We also include controls for household income and the number of persons in the household; income is included using a spline function where quintiles in the income distribution define the knots. In some specifications we also control for unobserved regional factors by including indicator variables to represent the nine census divisions in the United States. Finally, in every specification the standard errors of marginal effects are based on robust variance estimates that control for clustering of observations within states.

### *State-Level Analysis*

Ecological regressions are subject to bias, but they do hold two distinct advantages. First, we are able to examine a wider array of dependent variables than in our individual-level analysis. Second, we are able to construct a panel dataset that spans several decades; this permits us to control for unobserved state-level effects. In this study we examine the state-level association between 10- and 20-year lagged inequality and mortality for decade intervals from 1960 to 1990. We also examine this association for two leading causes of mortality, cardiovascular disease and malignant neoplasms, as well as for two specific causes that appear to be more strongly associated with inequality, homicides and accidents. The state-level mortality data are taken from the *Vital Statistics of the United States*. In each regression, we control for mean state income and the age-composition of state populations;<sup>4</sup> these data are calculated from decennial censuses and reported in the *Statistical Abstract of the United States*. In some regressions, we also control for unobserved state differences by including state fixed-effects.

## RESULTS

### *Individual-Level Analysis*

We report the estimated marginal effects of lagged inequality on health status for each year from 1995 to 1999; we run separate models using the state-level Gini for 1990, 1980, or 1970. As a result, we can test the significance of the Gini lagged at least five years (1995 CPS sample, 1990 Gini) to a maximum of

29 years (1999 CPS, 1970 Gini), as well as numerous other lengths of time in between. In Panel A of Table 1, we report the results of estimates that do not include controls for unobserved regional factors. Each cell shows the estimated marginal effect of a lagged Gini from a separate regression. The marginal effect of inequality is statistically significant at the 5 percent level in 6 of 15 regressions, significant at the 10 percent level in 5 regressions, and insignificant in 4 cases.<sup>5</sup> There is at best very weak evidence that the estimated marginal effects are largest when inequality is lagged approximately 15 to 19 years; this is broadly consistent with the evidence presented by Blakely et al. (2000), even though we include more individual characteristics as control variables in our analysis and we model household income as a spline function.<sup>6</sup>

These results suggest a potentially substantive adverse impact of inequality on health status. For example, a one standard deviation increase in lagged inequality implies *at most* a 0.9 percentage point increase in the probability that an individual reports fair or poor health status. This represents about a 7 percent increase in fair or poor health status, relative to the mean probability. It is unclear whether this should be considered an important or large impact, since a one standard deviation change in inequality may be unattainable through existing policy levers; on the other hand, mean state inequality increased by a little more than this amount from 1980 to 1990. Regardless, before we attribute this increase in the mean probability of fair or poor health status directly to state-level inequality, we next test the robustness of this finding against models that control for other regional factors. In particular, since both inequality and a number of health inputs (such as risk factors, lifestyles, and medical care access and quality) are known to vary across regions of the United States, we re-estimate our models after including indicator variables for the nine regional census divisions.

The results in Panel B of Table 1 tell a very different story. Once we include fixed division effects, the marginal effects of inequality are dramatically smaller; most have the wrong sign and nearly all are statistically insignificant. Only in the 1999 sample do we find that lagged inequality has a positive association with individual fair/poor health status, albeit one that is not statistically significant. In fact, the estimated effect of a 15- or 25-year lag is actually significantly and *negatively* associated with health status in the 1995 sample. We tested the null hypothesis that each of the point estimates in Panel B has a value equal to its counterpart in Panel A. In every year except 1999, we were able to reject this null hypothesis ( $p < 0.02$ ). We also tested the joint significance of the division dummies in each probit model reported in Panel B

Table 1: Estimated Effects of Lagged Income Inequality on Poor Health Status

| <i>Panel A: Models Control for Individual Characteristics, Mean State Income</i>                          |          |          |          |          |          |
|---|----------|----------|----------|----------|----------|
|   | 1995 CPS | 1996 CPS | 1997 CPS | 1998 CPS | 1999 CPS |
| 1990  | 0.179    | 0.305    | 0.280    | 0.228    | 0.351    |
| Gini  | (0.96)   | (2.01)   | (1.83)   | (2.10)   | (3.05)   |
| 1980  | 0.208    | 0.291    | 0.313    | 0.284    | 0.330    |
| Gini  | (1.13)   | (1.72)   | (1.81)   | (2.31)   | (2.58)   |
| 1970  | 0.098    | 0.186    | 0.235    | 0.192    | 0.236    |
| Gini  | (0.76)   | (1.39)   | (1.85)   | (1.83)   | (2.28)   |
| <i>Panel B: Models Control for Individual Characteristics, Mean State Income, Census Division Dummies</i> |          |          |          |          |          |
|   | 1995 CPS | 1996 CPS | 1997 CPS | 1998 CPS | 1999 CPS |
| 1990  | -0.162   | -0.106   | -0.104   | -0.090   | 0.157    |
| Gini  | (1.33)   | (0.73)   | (0.63)   | (0.93)   | (1.37)   |
| 1980  | -0.301   | -0.273   | -0.238   | -0.154   | 0.165    |
| Gini  | (2.23)   | (1.63)   | (1.14)   | (1.15)   | (1.20)   |
| 1970  | -0.263   | -0.240   | -0.231   | -0.166   | 0.174    |
| Gini  | (2.63)   | (1.65)   | (1.43)   | (1.38)   | (1.54)   |
| N   | 65,589   | 60,161   | 61,026   | 60,972   | 61,387   |

*Table notes:* Each cell contains the marginal effect of a lagged Gini coefficient estimated from separate probit models of the probability of reporting fair or poor health. Marginal effects were calculated by multiplying the probit coefficient by the value of the standard normal probability density function evaluated at the mean of all explanatory variables. Absolute values of t-statistics are reported in parentheses; the test statistic is based on the null hypothesis that the underlying coefficient is equal to zero. Standard errors are based on robust variance estimates that control for the clustering of observations within states. The individual characteristics used as explanatory variables were age, age squared, race, Hispanic ethnicity, education level, sex, residence in a metropolitan area and central city area, marital status, health insurance coverage status, household income (specified as a spline function using quintiles in the income distribution as the knots), and the number of persons in the household.

of Table 1 using Wald tests; in each of the 15 models, we were able to reject the null hypothesis that the coefficients on these division dummies were jointly equal to zero at levels of 0.00001 or better. Overall, these findings offer little support for the contention that exposure to income disparities causes lower health status. Instead, the association between inequality and health that was initially observed in Panel A is most likely biased upward by omitted controls for regional variations in smoking, diet, exercise, occupations, and medical care.

An additional claim made in the literature is that older adults are more susceptible to the adverse health consequences of prolonged exposure to inequality. Specifically, the results reported by Blakely et al. (2000) show a positive and significant association between income inequality and poor or fair health only for adults older than age 44. To test whether this finding holds up to the inclusion of additional individual-level covariates and geographic controls, we estimate our models replacing the Gini coefficient with two interaction terms. The first is the Gini coefficient multiplied by an indicator variable equal to one if the respondent is age 45 or older, and second term is the Gini coefficient multiplied by a dummy variable equal to one if the respondent is younger than age 45.<sup>7</sup> In these models, we found a pattern very similar to that seen in Table 1. Without controls for geographic region, the marginal effects of the Gini interaction terms were positive, and were statistically significant in many, but not all, cases. With the addition of division controls, the marginal effects of the Gini interaction terms were usually negative and statistically insignificant. The exceptions to this pattern were in the 1999 sample, in which the marginal effects remained positive, but insignificant, and in the 1995 sample, where the marginal effects of the 1980 and 1970 income inequality interaction terms were negative and significant. Regarding the size of the effect, the only evidence that we find that supports a stronger adverse effect for adults aged 45 and older comes from the aforementioned 1999 sample. For example, using the 1980 Gini in the 1999 sample, we estimate that inequality had a marginal effect of 0.18 for the older group and an effect of 0.14 for the younger group (as we note above, these effects were not statistically significant). However, we did observe some cases in which negative and statistically significant marginal effects of income inequality were larger in absolute value for the younger group. For example, in the 1995 sample, the 1980 Gini had a marginal effect of  $-0.334$  ( $t = 2.44$ ) for those younger than age 45, and a marginal effect of  $-0.283$  ( $t = 2.13$ ) for those aged 45 and older.

Since the inclusion of indicator variables for census division and state fixed-effects results in a model in which deviations from mean health are regressed on deviations from mean income inequality, it is important that a sizeable amount of variation in inequality remains once division effects (and similarly, state effects) are taken into account. As shown in appendix Table A2, the state-level Gini ranges from a minimum of 0.353 to a maximum of 0.489 in 1990, with slightly shorter ranges in the previous two decades. To test how much variation in this measure remains once division dummies are added to the model, we regressed the Gini coefficient on the set of division dummies



separately for each decade. The R-squared measures from these models were 0.44, 0.57, and 0.67 for 1990, 1980, and 1970 respectively, suggesting that one-third to more than one-half of the variation in the Gini remains once division effects are included. Taken together, these tests offer additional support for the model specification employed in our analysis. However, even if attenuation bias is present (from a reduction in signal-to-noise produced by deviations from the mean), this cannot account for the fact that most of our estimates in Panel B of Table 1 have the opposite sign of what would be predicted by the income inequality hypothesis. If anything, attenuation bias causes us to understate the association between inequality and *improved* health status.

Finally, we also tested the sensitivity of our key findings to the use of other regional controls, namely indicator variables for geographic regions in the United States (South, Midwest and West, relative to East), or an indicator variable for residence in the South. In these two sets of alternate specifications, we found the same pattern observed in Panel B of Table 1; 13 of 15 coefficient estimates are either insignificant or significant and negative (the exceptions are again in the 1999 sample, for the 1980 and 1990 Ginis). In addition, we were able to reject the null hypothesis that the region effects are jointly equal to zero in all specifications except the two noted above. Similarly, we tested the significance of the South indicator variable in all models; it was significantly different from zero in every year except 1999.

### *State-Level Analysis*

We report regression estimates of the effect of 10- and 20-year lagged inequality on state-level overall and cause-specific mortality rates per 100,000 population in Table 2. Descriptive statistics for mortality rates and lagged Gini coefficients are reported in Table A3. In Panel A of Table 2, we present the results of estimation without controls for state fixed-effects. Both the 10- and 20-year lagged Gini coefficients are significantly associated with higher mortality rates. These coefficients suggest that a one standard deviation increase in the inequality measure would lead to about a 1.9 percent increase in mean mortality over 10 or 20 years. Inequality is more strongly associated with homicides and accidents; for example, a one standard deviation change in the 10-year lagged Gini is associated with about a 50 percent increase in homicides and a 6 percent increase in accidents. However, no significant association is found for either cardiovascular disease or malignant neoplasms; in fact, the association has the wrong sign in three of the four models. This is particularly troubling for the income inequality hypothesis, since one of the

Table 2: Estimated Effects of Lagged Income Inequality on State Mortality Rates

*Panel A: Models Control for State Mean Income, Age Composition of State Population, and Decade Effects*

|                                       | <i>All Causes</i> | <i>Cardiovascular Disease</i> | <i>Malignant Neoplasms</i> | <i>Homicides</i> | <i>Accidents</i> |
|---------------------------------------|-------------------|-------------------------------|----------------------------|------------------|------------------|
| 10 year lagged Gini [ <i>n</i> = 192] | 436.4<br>(3.30)   | - 137.5<br>(1.37)             | - 58.8<br>(1.80)           | 81.9<br>(10.20)  | 85.6<br>(4.32)   |
| 20 year lagged Gini [ <i>n</i> = 144] | 385.2<br>(3.10)   | 8.6<br>(0.10)                 | - 38.4<br>(1.16)           | 77.2<br>(7.12)   | 99.8<br>(5.52)   |

*Panel B: Models Control for Mean Income, Age Composition of State Population, Decade Effects, and State Effects*

|                                       | <i>All Causes</i> | <i>Cardiovascular Disease</i> | <i>Malignant Neoplasms</i> | <i>Homicides</i> | <i>Accidents</i> |
|---------------------------------------|-------------------|-------------------------------|----------------------------|------------------|------------------|
| 10 year lagged Gini [ <i>n</i> = 192] | - 76.6<br>(0.53)  | - 198.1<br>(1.47)             | - 184.9<br>(5.16)          | 24.6<br>(3.85)   | 21.1<br>(1.20)   |
| 20 year lagged Gini [ <i>n</i> = 144] | - 58.0<br>(0.69)  | - 157.9<br>(2.03)             | - 162.9<br>(4.00)          | 8.48<br>(0.90)   | - 1.45<br>(0.08) |

*Table Notes:* Each cell contains the regression coefficient estimate for the lagged Gini coefficient. Absolute values of t-statistics are reported in parentheses (based on heteroskedastic-consistent standard errors); the test statistic is based on the null hypothesis that the coefficient is equal to zero.

major pathways conjectured to link inequality to health outcomes is through psychosocial stress and heart disease.

Once again, we get very different results once we account for unobserved fixed factors.<sup>8</sup> The results in Panel B of Table 2 demonstrate that exposure to inequality is unrelated to overall mortality rates. Now the coefficient on inequality is negative, though not significant; in addition, we reject the null hypothesis that the inequality coefficients in models of all-cause mortality in Panel B of Table 2 are identical to those in Panel A. Further, even if we take the upper bound of the 95 percent confidence interval for the Panel B estimates, a one standard deviation increase in lagged inequality is associated with at most an increase of about 0.1 percent in all cause mortality. Prior exposure is significantly associated with lower mortality rates for cardiovascular disease and malignant neoplasms in three of four specifications, while homicide mortality is significantly associated only with 10-year lagged with inequality.

We do not wish to put too much weight on these findings, given the pitfalls of ecological inference.<sup>9</sup> But the results for all-cause mortality do accord with our findings in the earlier individual analysis: once we control for unobserved state effects, there is no evidence of an association with exposure to income inequality. It is worth noting that the results for cardiovascular disease run counter to the notion that income inequality causes stress, and in turn cardiovascular disease (Wilkinson 1996).

We have also estimated these state-level regressions without mean income as control variable, as well as with additional state-level socio-economic control variables (percent black, percent with college or high school education, and percent urban); the overall pattern of results from these regressions is similar to that reported in Table 2.

Again, because we control for state fixed-effects, it is important to demonstrate that there remains sufficient variation in our lagged inequality variable in this analysis. A regression of either the 10-year or 20-year lagged Gini on state dummies yields R-squared measures of 0.60 and 0.67, respectively. This suggests that at least one-third of the variance in the lagged Gini measures remains after controlling for state fixed-effects. Further support for our approach is found in the fact that several of the estimated coefficients on lagged inequality are in fact statistically significant, albeit sometimes with the opposite sign of that predicted by the income inequality hypothesis. Finally, even if attenuation bias is present, it only causes us to understate the association between inequality and *reductions* in all-cause mortality.

## DISCUSSION

Our analysis offers little support for the hypothesis that prolonged exposure to income inequality holds adverse consequences for individual and population health. We found a positive and significant association between lagged income inequality and health status until we controlled for differences in geographic region—represented in our models by a set of census division indicator variables in the individual-level analysis and state fixed-effects in the state-level analysis. It is crucial to control for such factors, since important determinants of both health outcomes and income inequality vary across regions of the United States. Failure to control for these regional influences will yield biased estimates of the association between inequality and health, as has been reported by so many previous studies.

This study should give pause to those that would make policy recommendations based on previous findings; however, our own results must be interpreted with caution, as well. In general, we do not find that inequality is significantly associated with health status or all-cause mortality in our preferred models, but the standard errors on the estimated coefficients for the inequality variables are often large enough that we also cannot reject the null hypothesis that inequality has some modest effect on health status (albeit positive or negative). Further, we have followed much of the prior literature in focusing on state-level inequality; perhaps the kind of inequality that matters for health is that which occurs at the regional level, or alternatively at the neighborhood level, or within a peer group. Consequently, we cannot rule out the possibility that inequality defined at some level or across some groups may somehow impact health, but we doubt the existence of a significant connection between state-level inequality and either individual health status or state mortality rates.

## APPENDIX

Table A1: Means and Standard Deviations

|                     | <i>1995 CPS</i>    | <i>1996 CPS</i>    | <i>1997 CPS</i>    | <i>1998 CPS</i>    | <i>1999 CPS</i>    |
|---------------------|--------------------|--------------------|--------------------|--------------------|--------------------|
| Poor or fair health | 0.143<br>(0.350)   | 0.140<br>(0.347)   | 0.141<br>(0.349)   | 0.136<br>(0.343)   | 0.138<br>(0.345)   |
| Household income    | 48,128<br>(36,092) | 50,888<br>(47,927) | 51,967<br>(49,772) | 54,169<br>(52,430) | 55,498<br>(52,575) |
| Mean state income   | 40,868<br>(4,474)  | 43,299<br>(5,165)  | 44,277<br>(5,323)  | 46,005<br>(5,522)  | 47,132<br>(5,501)  |
| <i>N</i>            | 65,589             | 60,161             | 61,026             | 60,972             | 61,387             |

Source: Authors' calculations using various years of the March Current Population Survey.

Table A2: Means and Standard Deviations of Gini Coefficient (across 50 U.S. States and the District of Columbia)

|                    | <i>1990 Gini</i> | <i>1980 Gini</i> | <i>1970 Gini</i> |
|--------------------|------------------|------------------|------------------|
| Mean               | 0.400            | 0.358            | 0.358            |
| Standard deviation | 0.026            | 0.022            | 0.027            |
| Min                | 0.353            | 0.316            | 0.313            |
| Max                | 0.489            | 0.434            | 0.425            |

Source: U.S. Census Bureau ([www.census.gov/hhes/income/histinc/state/state4.html](http://www.census.gov/hhes/income/histinc/state/state4.html)).

Table A3: Means and Standard Deviations (across 48 U.S. States)

|   | 1960–1990, by Decade | 1970–1990, by Decade |
|---|----------------------|----------------------|
| <i>Mortality Rates per 100,000</i>        |                      |                      |
| All causes                                | 900.8<br>(108.9)     | 888.9<br>(109.7)     |
| Cardiovascular disease                    | 444.5<br>(89.9)      | 429.0<br>(85.1)      |
| Malignant neoplasms                       | 168.5<br>(37.3)      | 178.9<br>(34.3)      |
| Homicides                                 | 6.7<br>(4.1)         | 7.4<br>(4.1)         |
| Accidents                                 | 52.5<br>(14.1)       | 50.7<br>(14.3)       |
| <i>Gini Coefficient for Family Income</i> |                      |                      |
| 10-Year Lag                               | .371<br>(.040)       |                      |
| 20-Year Lag                               |                      | .376<br>(.043)       |
| <i>N</i>                                  | 192                  | 144                  |

*Source:* Mortality rates are calculated by the authors from *Vital Statistics of the United States*, income inequality data are from Al-Samarrie and Miller (1967) and the U.S. Census Bureau ([www.census.gov/hhes/income/histinc/state/state4.html](http://www.census.gov/hhes/income/histinc/state/state4.html)).

## NOTES

1. For a recent review of the literature, see Wagstaff and van Doorslaer (2000).
2. The 1995 CPS was the first year in which a question pertaining to health status, our dependent variable, was included in the survey.
3. Individual respondents rate health status on a five point scale (excellent, very good, good, fair, or poor).
4. Several previous ecological studies in this literature have first age-adjusted state-level mortality, then performed multivariate regression using non-age-adjusted independent variables; however, this method is inappropriate since income inequality and mean income each co-vary with the age composition of state populations (Rosenbaum and Rubin 1984).
5. Some caution is in order here. The CPS sample rotation is such that one-half of the respondents in one year are included in the next year; therefore, Table 1 really contains at best three fully independent samples. Further, every regression shares the same set of individual explanatory variables; only the year in which the Gini is measured changes across specifications. Nevertheless, we report results for all permutations of the five CPS years and the three Ginis for completeness and to allow an evaluation of the claim made by Blakely et al. (2000) that inequality is most strongly associated with health status when lagged 15 years.

6. Our specification controls for several individual-level characteristics. One of these, insurance coverage, may be in part determined itself by health status, since sicker individuals have a greater demand for private insurance. Others argue that inequality is a determinant of educational attainment (e.g., Blakely, Lochner, and Kawachi 2002), so that by including controls for education, we may understate the effect of inequality on health. Consequently, we have also estimated our models omitting either or both controls for insurance coverage and education; the basic pattern of results is unchanged.
7. We conducted an alternate test of this hypothesis by separating the sample into groups of individuals aged 45 and older, and younger than 45, thus allowing the marginal effects of all explanatory variables to vary by age. We do not find positive and significant effects of inequality for either group, except for the older group in the 1999 sample ( $p < .10$ ).
8. We tested the joint significance of the state dummies in each regression model reported in Table 2, and we were able to reject the null hypothesis that all coefficients were equal to zero in every case.
9. For example, in contrast to these ecological associations, Meara (1999) finds no significant relationship between state income inequality and infant health.

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