

Socioeconomic Inequality in Voting Participation and Self-Rated Health

ABSTRACT

Objectives. This study tested the hypothesis that disparities in political participation across socioeconomic status affect health. Specifically, the association of voting inequality at the state level with individual self-rated health was examined.

Methods. A multilevel study of 279066 respondents to the Current Population Survey (CPS) was conducted. State-level inequality in voting turnout by socioeconomic status (family income and educational attainment) was derived from November CPS data for 1990, 1992, 1994, and 1996.

Results. Individuals living in the states with the highest voting inequality had an odds ratio of fair/poor self-rated health of 1.43 (95% confidence interval [CI]=1.22, 1.68) compared with individuals living in the states with the lowest voting inequality. This odds ratio decreased to 1.34 (95% CI=1.14, 1.56) when state income inequality was added and to 1.27 (95% CI=1.10, 1.45) when state median income was included. The deleterious effect of low individual household income on self-rated health was most pronounced among states with the greatest voting and income inequality.

Conclusions. Socioeconomic inequality in political participation (as measured by voter turnout) is associated with poor self-rated health, independently of both income inequality and state median household income. (*Am J Public Health*. 2001;91:99–104)

Tony A. Blakely, MBChB, MPH, Bruce P. Kennedy, EdD, and Ichiro Kawachi, PhD

Higher income inequality at the state level in the United States has been associated with increased mortality in ecologic studies^{1,2} and in multilevel studies of self-rated health.^{3,4} Income inequality has also been linked with mortality at the metropolitan statistical area level⁵ and with self-rated health at the county and census tract levels.⁴ There have also been studies that failed to find an association between income inequality and health.^{6,7}

There are at least 3 plausible mechanisms linking income distribution to health: variations in an individual's access to life opportunities and material resources (e.g., health care, education), the degree of social cohesion or differential investment in social capital, and possible direct psychosocial processes related to relative perceptions of position in the socioeconomic hierarchy.^{8,9} The first 2 of these pathways are in turn likely to be mediated through political processes. We have argued elsewhere^{8–10} that a wide gap between the assets of the rich and the poor leads to a corresponding polarization of the political interests of groups in society, which is reflected in diminishing levels of social trust and reduced spending on social goods that benefit all (e.g., basic education, health care). In other words, we posit a direct relationship between growing inequality and disparities in political participation across socioeconomic groups, which translates into policies that are detrimental to health.

An independent line of evidence in political science suggests that socioeconomic inequalities in political participation lead to state policies that harm the poor. Thus, Hill and Leighley reported that lower representation of poorer voters at the state level in 1986 (both of itself and relative to high-income voters) was associated with less generous welfare spending by states.¹¹ Subsequently, Hill et al. replicated this finding for 1978 to 1990,¹² although the association weakened over the 1980s, possibly owing to constrained state budgets during the period of "New Federalism." Aside from the act of voting, monetary donations to political campaigns dispropor-

tionately come from people of higher socioeconomic status,¹³ and this in turn may cause policy to be skewed to the interests of these groups.¹⁴ Given the foregoing, then, it is plausible that unequal political participation at the state level is also associated with worse average health status.

A related explanation is that political participation is both a measure of civic engagement generally and a proxy measure of the broader concept of social capital.¹⁵ Social capital—as indexed by density of membership in associations, levels of interpersonal trust, and strength of norms of mutual aid and reciprocity^{15,16}—has been associated with mortality and self-rated health in the United States.^{10,17} Kawachi and Kennedy have reported a correlation between levels of social capital at the state level and overall voter turnout at election time.¹⁸ Kawachi et al. have also reported a correlation at the state level of –0.51 between a composite index of women's political participation and female mortality rates.¹⁹

The primary objective of this study was to examine the relation between socioeconomic inequality in political participation at the state level in the United States and individual self-rated health. States were selected as the relevant unit of analysis, because legislation, taxation policies, and welfare programs vary between states. To measure socioeconomic inequality in political participation, we used data on voting; it would have been ideal to also use data on other forms of inequality of political par-

At the time of the study, all of the authors were with the Department of Health and Social Behavior, Harvard School of Public Health, and Harvard Center for Society and Health, Boston, Mass. Tony A. Blakely is now with the Department of Public Health, Wellington School of Medicine, University of Otago, Wellington, New Zealand.

Requests for reprints should be sent to Tony A. Blakely, MBChB, MPH, Department of Public Health, Wellington School of Medicine, University of Otago, PO Box 7343, Wellington, New Zealand (e-mail: tblakely@wnmeds.ac.nz).

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ticipation (e.g., campaign financing), but the necessary data were not available. A secondary objective was to examine how income inequality and voting inequality—2 separate measures of “socioeconomic inequality”—were independently and jointly associated with health. Although our focus was on inequality per se, we also examined the association of overall voter turnout with self-rated health.

Methods

Individual-Level Sociodemographic and Health Data

For data on individual-level covariates, we combined the March Current Population Survey (CPS) Supplements for 1995 and 1997.²⁰ Each household is sampled by the CPS for 2 years, so 1996 data were discarded to avoid count-

ing individuals twice. Since 1995, the March Supplement has collected information on self-rated health in 5 categories (excellent, very good, good, fair, and poor) by means of direct questioning of each adult in the household and of parental proxy for children. A review of 27 studies has shown that this simple measure of self-rated health has strong predictive validity for mortality, independent of other physiologic, behavioral, and psychosocial factors.²¹ The self-rated health variable was dichotomized as fair and poor vs the 3 other responses. We adjusted household income for household size with a standard “equivalization” procedure (i.e., dividing the household income by the square root of the number of people in that household) and then created a 9-level categorical variable (see Table 1 for categories). Remaining individual-level variables were sex, age (0–14, 15–24, 25–34, 35–44, 45–54, 55–64, 65–74, 75 years and older), and race (White, Black, other).

State-Level Voting Turnout Inequality

Voting turnout inequality measures were calculated from November CPS Voting Supplements for 1990, 1992, 1994, and 1996. For each of these 4 samples, we excluded all individuals aged under 25 years, because educational attainment (used to measure socioeconomic inequality in voting) would often not be completed at younger ages. We also excluded those ineligible to vote and individuals who did not respond either “yes” or “no” to the question “Did you vote in the election held on ———?” (approximately 7% each year). This resulted in 84432, 82047, 78786, and 68929 observations in each of the 4 samples, respectively.

For each state, we calculated a relative index of inequality (RII)^{22,23} for voting by family income. Briefly, the population in each state was assigned to categories of family income, and each category was assigned a midpoint on a cumulative proportion distribution. For example, if 40% of individuals in a state had a family income less than \$25 000, and the category with a family income of \$25 000 to \$30 000 comprised 10% of individuals, then the midpoint on a cumulative proportion distribution for this category would be 0.45 (0.40 + [0.10/2]).

The proportion of individuals voting in each category of family income was then regressed (by ordinary least squares regression) on the midpoint for each category, generating an intercept and slope for each state. This method for calculating the RII assumes that the underlying relationship is linear; scatter plots supported this assumption. The intercept from the regression model for each state is the probability that the individual with the lowest family income in each state voted. Because the midpoints range from 0 to 1, the slope from the regression model is the absolute difference in the probability of voting between the individuals with the lowest and highest family incomes. The RII was calculated by summing the intercept and the slope and then dividing by the intercept—that is, it is the hypothetical relative difference in probability of voting between the individuals with highest and lowest family incomes. Following this procedure, 4 RIIs (1 for each year) of voting inequality by family income were calculated for each of the 50 states. RIIs of voting inequality by level of educational attainment were similarly calculated for each of the 50 states.

There was little evidence that the income RIIs were measuring a different construct from the education RIIs, or that the presidential and nonpresidential election year RIIs were measuring a different construct (correlation coefficients available from the authors on request). Therefore, we averaged the RIIs into a single summary measure. We then assigned the 50 states to 4 categories of voting turnout inequality, using the mean of this average RII

TABLE 1—Distribution of March Current Population Survey Supplement Respondents (1995 and 1997) by Demographic Factors, Equivalized Household Income, and Socioeconomic Inequality in Voting, and Their Associated Odds Ratios of Fair/Poor Health^a

	n (total = 279 066)	% Fair/Poor Health (Nonweighted)	Odds Ratio (95% CI)
Sex			
Female	145 001	12.7	0.99 (0.96, 1.01)
Male ^b	134 065	10.4	1.00
Age, y			
0–14	65 371	3.0	0.41 (0.38, 0.43)
15–24	37 211	3.8	0.59 (0.55, 0.63)
25–34 ^b	41 891	5.6	1.00
35–44	44 589	8.9	1.87 (1.77, 1.97)
45–54	33 672	14.4	3.86 (3.65, 4.08)
55–64	22 514	22.6	5.88 (5.56, 6.22)
65–74	19 364	32.9	8.05 (7.61, 8.50)
≥75	14 454	44.0	11.4 (10.7, 12.1)
Race			
Other	16 453	10.2	1.19 (1.11, 1.27)
Black	27 259	16.4	1.55 (1.49, 1.61)
White ^b	235 354	11.1	1.00
Equivalized household income, \$			
<5 000	18 927	18.1	7.67 (7.15, 8.23)
5 000–9 999	33 917	23.5	7.08 (6.65, 7.54)
10 000–14 999	37 400	16.9	4.27 (4.01, 4.54)
15 000–19 999	36 007	12.1	3.15 (2.95, 3.36)
20 000–24 999	32 103	9.1	2.34 (2.19, 2.50)
25 000–29 999	27 767	7.6	1.96 (1.82, 2.10)
30 000–39 999	38 847	6.2	1.52 (1.42, 1.63)
40 000–49 999	22 593	5.6	1.31 (1.21, 1.41)
≥50 000 ^b	31 505	5.0	1.00
Voting inequality			
High (9 states)	50 607	14.1	1.43 (1.22, 1.68)
Medium-high (12 states)	48 195	11.7	1.12 (0.96, 1.31)
Medium-low (20 states)	129 641	11.1	1.00 (0.87, 1.15)
Low ^b (9 states)	50 623	10.2	1.00

Note. CI = confidence interval.

^aThe odds ratios are from a weighted logistic regression model for fair/poor self-rated health that included all variables in this table simultaneously, with a random intercept at the state level.

^bReference category.

and the mean ± 1 standard deviation as cut-points. In addition to the RIIs, we also calculated the average voter turnout for each state, averaged over the 4 years.

State-Level Income Inequality and Median Income

Gini coefficients of equivalized household income for each state, adjusted for federal and state income and payroll taxes as well as for cash or near-cash benefits including food stamps, the earned income tax credit, and school lunches, were obtained from the Luxembourg Income Study (Timothy Smeeding, project director, oral communication, September 1996). We calculated the Gini values by using pooled 1991, 1992, and 1993 March CPS data. The Gini coefficient ranges theoretically from 0 (absolute equality) to 1.0 (absolute inequality in the distribution of income) (see Kawachi and Kennedy²⁴ and Ryscavage²⁵ for details on calculation). State-level median income was calculated from pooled 1991 and 1993 March CPS data.²⁰

Multilevel Data Analysis

The state-level measures of voting inequality, income inequality, and median income were modeled with the individual-level data from the 1995 and 1997 March CPS data set. We fitted logistic random effects models by using Proc Glimmix in SAS.²⁶ All models included a random error term at both the individual and state levels and were controlled for age, sex, race, and equivalized household income at the individual level. The March CPS Supplement is a complex survey, designed to give national and state-level estimates of income and other sociodemographic measures.²⁷ The publicly available CPS data do not include sufficient sampling information to conduct an analysis allowing for design effects, but individual and household weights are provided. Use of sampling weights should give accurate point estimates for regression coefficients, but the standard errors may still not be correct.²⁸ The CPS weights were used in all regression analyses; results for unweighted analyses (not presented) were not substantially different.

To investigate the combined effects of voting and income inequality, we reassigned states to 5 categories on the basis of a cross-classification of voting and income inequality. Possible cross-level effect modification²⁹ of the combined voting and income inequality variable with demographic variables was investigated via separate models for sex, age group, and race. For individuals with an equivalized household income greater than \$5000 (93.2% of the sample), the association between the log-arithm of household income and self-rated health was linear in the logit. Therefore, to determine cross-level effect modification of state-

level inequality and household income, we ran 1 model restricted to subjects with household incomes greater than \$5000, with variables for log household income, dummy variables for the combined voting and income inequality variable, and interaction products of these variables. The predicted probabilities of fair/poor health from the model were then used to assess the presence of effect modification (in terms of departure from risk additivity).

Results

The average *national* voter turnout among those older than 25 years was 56.5%, 72.8%, 56.4%, and 66.9% for 1990, 1992, 1994, and 1996, respectively. When voter turnout was calculated separately by state and then averaged over the 4 years, the average state-level voter turnout was 63.0% (SD=6.1%), with a range of 48.9% (West Virginia) to 74.4% (Minnesota).

Both family income and educational attainment were strongly associated with voting. The RIIs for the *national* CPS Voting Supplement samples (1990, 1992, 1994, and 1996) for education were 2.02, 1.94, 2.29, and 2.04 and for family income were 1.78, 1.74, 2.04, and 1.78, respectively. That is, using the 1990 sample as an example, the probability of voting among individuals with the highest education in the United States was 2.02 times greater than that for individuals with the lowest education. The *state-level* average RIIs ranged from 1.50 (Illinois) to 3.09 (West Virginia). The average voter turnout and average RIIs by state were highly correlated ($r=-0.75$)—a reflection of the fact that the proportion of people voting is bounded between 0 and 1 and that as the average voter turnout increases, the RII tends toward 1. That is, the average voter turnout and average RII are statistically related measures.

Table 1 shows the odds ratios of self-rated fair/poor health for individuals living in states with varying degrees of voting inequality, with sex, age, race, and equivalized household income controlled for at the individual level. Individuals living in states with high voting inequality had an odds ratio of 1.43 (95% confidence interval [CI]=1.22, 1.68) for fair/poor self-rated health, compared with individuals in states with low voting inequality. The corresponding odds ratio for individuals in states with medium-high voting inequality was 1.12 (95% CI=0.96, 1.31), and for individuals in states with medium-low levels of voting inequality it was 1.00 (95% CI=0.87, 1.15). Introducing income inequality (Table 2, first column) into the model reduced the odds ratio for individuals living in states with high voting inequality to 1.34 (95% CI=1.14, 1.56). Greater income inequality was associated with fair/poor health independently of voting inequality.

The effect of adding state median income (Table 2, second column) was to reduce the odds ratio for both the states with the highest voting inequality and those with the highest income inequality, although the odds ratios for the most unequal states still remained elevated.

Table 2 also presents results for models that include the average voter turnout for each state, rather than the voting inequality. The odds ratios by level of average voter turnout are approximately a third greater than the corresponding values for voting inequality. Including both average voter turnout and voting inequality (but not income inequality or median income) in the same model resulted in an odds ratio of 1.46 (95% CI=1.20, 1.79) for the states with low voter turnout and 1.14 (95% CI=0.95, 1.36) for the states with high voting inequality. However, because of the mathematical interdependence of these 2 measures, 5 of the 9 states with high voting inequality were also low-voter-turnout states, and 6 of 9 states with low voting inequality were also high-voter-turnout states. Thus, these results should be interpreted with caution.

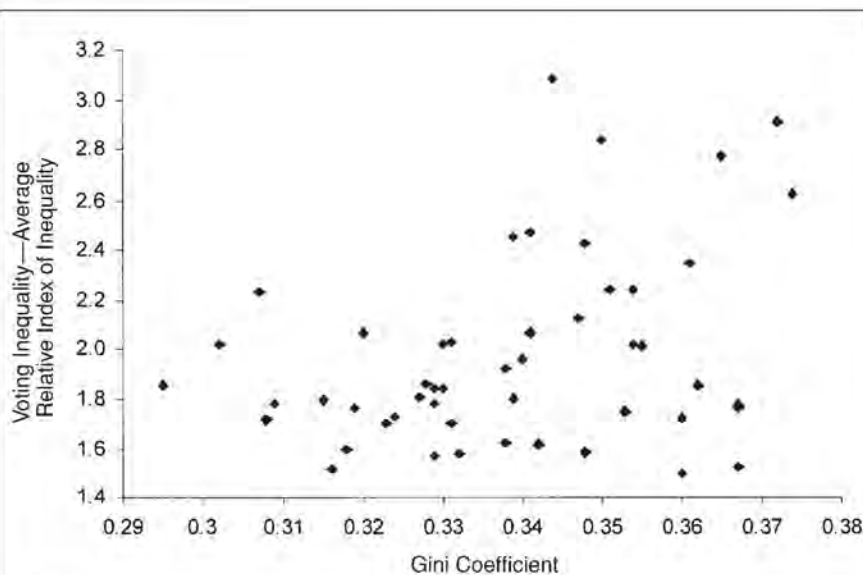
The plot of the average RII vs the Gini coefficient (Figure 1) suggested a J-shaped association: there were no states with a low Gini coefficient *and* high levels of voting inequality. The combined effect of voting inequality and income inequality was investigated by forming 5 groups of states: 9 states with high voting inequality and above-average income inequality (West Virginia, Georgia, Virginia, Kentucky, Texas, South Carolina, North Carolina, Tennessee, and Alabama; top right of Figure 1); 7 states with medium-high voting inequality and above-average income inequality (Arkansas, Oklahoma, Colorado, Nevada, New Hampshire, Arizona, and Idaho; middle right of Figure 1); 10 states with medium-low or low voting inequality and above-average income inequality (California, Florida, Illinois, Louisiana, Massachusetts, Michigan, Mississippi, Missouri, New Jersey, and New York; bottom right of Figure 1); 5 states with medium-high voting inequality and below-average income inequality (Delaware, Maryland, New Mexico, Ohio, and Vermont; middle left of Figure 1); and the remaining 19 states with medium-low or low voting inequality and below-average income inequality (bottom left of Figure 1). Individuals living in 1 of the 9 states with high voting inequality *and* above-average income inequality had an odds ratio of self-rated fair/poor health of 1.54 (95% CI=1.36, 1.75), compared with individuals in the 19 states with low voting inequality and low income inequality (Table 3). The odds ratios for the 3 other nonreference categories of states were all 1.21. Including state median income reduced the excess odds ratio for states with high voting inequality and high income inequality by 25% (from 1.54 to 1.40), but it had

TABLE 2—Odds Ratios (95% Confidence Intervals) of Fair/Poor Health for Categories of State-Level Socioeconomic Inequality in Voting, Average Voter Turnout, Income Inequality, and Median Household Income^a

State-Level Variable (No. of States)	Voting and Income Inequality	Voting and Income Inequality and Median Household Income	Average Voter Turnout	Average Voter Turnout and Income Inequality	Average Voter Turnout, Income Inequality, Median Household Income
Voting inequality					
High (9)	1.34 (1.14, 1.56)	1.27 (1.10, 1.45)			
Medium-high (12)	1.11 (0.96, 1.30)	1.11 (0.97, 1.26)			
Medium-low (20)	1.03 (0.90, 1.18)	1.04 (0.93, 1.17)			
Low (9) ^b	1.00	1.00			
Income inequality					
High (10)	1.27 (1.09, 1.49)	1.18 (1.03, 1.36)		1.23 (1.07, 1.42)	1.17 (1.03, 1.34)
Medium-high (16)	1.25 (1.07, 1.46)	1.23 (1.08, 1.41)		1.25 (1.09, 1.42)	1.24 (1.10, 1.41)
Medium-low (15)	1.14 (0.99, 1.32)	1.13 (1.00, 1.28)		1.14 (1.00, 1.31)	1.13 (0.99, 1.28)
Low (9) ^b	1.00	1.00		1.00	1.00
Median household income					
Low (8)		1.31 (1.13, 1.51)			1.22 (1.04, 1.44)
Medium-low (20)		1.03 (0.92, 1.17)			1.04 (0.93, 1.17)
Medium-high (14)		1.05 (0.93, 1.19)			1.10 (0.96, 1.26)
High (8) ^b		1.00			1.00
Average voter turnout					
Low (7)			1.62 (1.39, 1.90)	1.48 (1.26, 1.75)	1.38 (1.16, 1.64)
Medium-low (19)			1.19 (1.05, 1.36)	1.14 (1.00, 1.30)	1.17 (1.03, 1.32)
Medium-high (15)			1.04 (0.90, 1.19)	1.02 (0.89, 1.16)	1.02 (0.90, 1.17)
High (9) ^b			1.00	1.00	1.00

^aThe odds ratios are from weighted logistic regression models for fair/poor self-rated health that included the variables in each column and individual-level sex, age, race, and equivalized household income. A random intercept was specified at the state level.

^bReference category.

**FIGURE 1—Plot of socioeconomic inequality in voting by income inequality for the 50 states.**

no substantive effect for the 3 other nonreference categories of states.

We next examined potential cross-level effect modification by sex, race, and age by stratifying models according to these demographic variables. There was effect modification by age group (Table 3), such that the effect of inequality was greater for older age groups (>45 years). For those aged 45 to 64 years, liv-

ing in the 9 states with high voting inequality and high income inequality conferred an odds ratio of 1.71 (95% CI=1.42, 2.05). The corresponding odds ratio for individuals 65 years and older was 1.65 (95% CI=1.42, 1.90). Regarding race, the odds ratios for high combined voting inequality and income inequality were not notably elevated among Blacks (Table 3; results for "other" not shown owing to small

numbers). Thus, living in more egalitarian states was associated with better self-rated health for Whites, but only marginally so for Blacks. No meaningful effect modification was apparent by sex.

For the 93.2% of the total sample with an equivalized household income greater than \$5000, we ran 1 model with dummy variables for the 5 groups of states formed by combining voting and income inequality, a variable for log household income, and interaction products of these variables. On a multiplicative scale, the "slope" of the association of individual household income with self-rated health was steepest within the 9 states with both high voting and high income inequality, although the 95% confidence intervals overlapped between the 5 groups of states (results available from the authors on request). That is, among states with the most inequality, the odds ratio of fair/poor self-rated health was greater than for the 4 other groups of states for each unit decrease in household income. Predicted values from the model for reference individuals (male, White, 25–34 years) with household incomes of \$5000 and \$50000 were used to assess effect modification on an additive scale. On the basis of these calculations, among White men aged 25 to 34 years living in states with the highest inequality, fair/poor self-rated health was reported by 2.4% of those with an equivalized household income of \$50 000 and by 19.2% of those with an equivalized household income of \$5000. That is a risk difference of 16.8 percentage points. By contrast,

TABLE 3—Odds Ratios (95% Confidence Intervals) of Fair/Poor Health for 5 Groups of States Cross-Classified by Voting Inequality and Income Inequality, by Age Group^a

Group	Voting Inequality (RII)	Income Inequality (Gini)	
		Low	High
All ages, all races	Low	1.00	1.21 (1.08, 1.37)
	Medium-high	1.21 (1.02, 1.43)	1.21 (1.05, 1.40)
	High	NA	1.54 (1.36, 1.75)
Age group, y 0–14 ^b	Low	1.00	1.16 (0.95, 1.42)
	Medium-high	1.29 (0.96, 1.73)	1.23 (0.94, 1.61)
	High	NA	1.13 (0.91, 1.41)
15–24 ^b	Low	1.00	1.21 (0.93, 1.57)
	Medium-high	1.15 (0.77, 1.71)	1.33 (0.94, 1.86)
	High	NA	1.23 (0.93, 1.63)
25–44 ^b	Low	1.00	1.17 (1.02, 1.35)
	Medium-high	1.27 (1.04, 1.56)	1.07 (0.90, 1.28)
	High	NA	1.35 (1.16, 1.56)
45–64 ^b	Low	1.00	1.31 (1.09, 1.56)
	Medium-high	1.15 (0.90, 1.48)	1.21 (0.97, 1.50)
	High	NA	1.71 (1.42, 2.05)
≥65 ^b	Low	1.00	1.10 (0.96, 1.26)
	Medium-high	1.17 (0.95, 1.42)	1.28 (1.08, 1.52)
	High	NA	1.65 (1.42, 1.90)
Race White ^c	Low	1.00	1.22 (1.08, 1.38)
	Medium-high	1.21 (1.02, 1.43)	1.23 (1.06, 1.42)
	High	NA	1.60 (1.41, 1.82)
Black ^c	Low	1.00	1.01 (0.82, 1.24)
	Medium-high	1.07 (0.80, 1.43)	1.03 (0.76, 1.39)
	High	NA	1.11 (0.90, 1.38)

Note. RII = relative index of inequality.

^aEach model includes sex, age, race (if not stratified by same), and equalized household income at the individual level. No states had low income inequality and high voting inequality, hence this cell is not applicable (NA). Each panel of results is taken from a separate, but similarly specified, logistic regression model, with a random intercept at the state level.

^bIncludes both sexes and all races.

^cIncludes both sexes and all ages.

the same risk difference was predicted to be between 11 and 13 percentage points for the 4 other groups of more equal states. Thus, on an additive scale, the model suggests a 30% to 50% stronger association of low individual income and poor self-rated health among the most unequal states.

Discussion

The first key finding of our study was that there was no overall association between income inequality and voting inequality as hypothesized, although there were no states with low income inequality that also had high inequality in political participation (Figure 1). A priori, we expected a closer association of these 2 measures of inequality. A possible reason for the lack of association was that the RIIs used to generate the average voting inequality measure were inaccurate. However, each of the 8 component RIIs was based on 68 000 to 85 000 individual observations, and averaging the RIIs over the 4 election cycles should have mitigated

against random error. More important, voting inequality is only 1 aspect of political participation. Other political activities that may vary by socioeconomic status include volunteering in political campaigns, contacting officials, organizing protests, running for office, and, of increasing importance, donating money.^{13,30}

A second key finding of our study was that individuals living in states with high voting inequality, particularly the highest category, had increased odds of self-rated fair/poor health compared with individuals living in states with lower voting inequality. When voting and income inequality are combined into 1 measure of inequality, people living in the 9 most unequal states had 54% greater odds of fair/poor self-rated health than people living in the 19 most egalitarian states. This excess odds of fair/poor self-rated health for people living in the most unequal states was greatest among 45- to 64-year-olds (odds ratio [OR] = 1.71) and those 65 years and older (OR = 1.65). Interestingly, there appeared to be little association of increasing state-level inequality with self-rated health among Blacks

(Table 3), although the confidence intervals were wide.

The relationship between lower individual income and poorer health is well established.³¹ A possible psychosocial mechanism to explain this association focuses on the adverse physiologic consequences of being lower in the socioeconomic hierarchy.³² It has been suggested that the same psychosocial mechanism may also underpin the association of income inequality with health.^{8(p xix)} Here, it is thought that increasing the overall extent of inequality in a society amplifies the underlying psychosocial comparisons up and down the income ladder and, conversely, decreasing inequality dampens such comparisons—that is, there is a cross-level effect modification.²⁹ Our findings partly support such a hypothesis: lower individual-level household income was more strongly associated with poorer self-rated health in the most unequal states, both in multiplicative and in additive terms. In additive terms, people living in the 9 most unequal states (compared with people living in the remaining states) experienced a 30% to 50% greater risk of fair/poor self-rated health for each unit decrease in household income.

Our primary focus in this study was on inequality—whether in the form of income inequality or voting inequality. However, the absolute level of voter turnout in each state was a more powerful predictor of fair/poor self-rated health than voting inequality (Table 2). On the other hand, overall voter turnout is highly (and necessarily) correlated with voting inequality.³³ Also, overall voter turnout is measured with greater precision than voting inequality—voting inequality requires multiple measures of voter turnout by strata of socioeconomic status, and the measure of socioeconomic status will also be subject to misclassification. Thus, if both overall voter turnout and voting inequality are equally important, greater imprecision in the measurement of voting inequality will result in absolute voter turnout appearing to dominate in statistical models.

We had no data on individual-level behavioral and biological variables that may be confounders or intermediary variables between inequality and health. On the other hand, a multilevel study of income inequality and self-rated health found that including such variables did not substantively alter the association of income inequality with self-rated health.³ A possible weakness of this study is the use of a single summary measure of self-rated health, but this simple measure has been shown to be predictive of mortality.²¹

We measured voting inequality over a period up to 7 years before the assessment of self-rated health. It is unlikely that any effect on health of inequality in political participation would be instantaneous. Instead, we would ex-

pect a lag period during the course of which inequality in political participation would “get under the skin” via policies and other mechanisms.²⁹ Blakely et al. specifically examined possible lag times in the association of income inequality with self-rated health.³⁴ There was a suggestion that the association among people 45 years and older was strongest for income inequality measured 15 years before self-rated health and weakest for income inequality measured contemporaneously. However, rankings of states by inequality do not change substantially over a 10-year period, limiting the ability of such analyses to determine time lags.

Earlier in this report, we offered 2 broad possible explanations for the association of political participation with health: inequality in political participation skews subsequent policy, and the association of political participation with health is a proxy for the more general association of social capital with health. These 2 explanations are not mutually exclusive. In this report, we have more specifically examined the inequality explanation and found evidence to support such an association. However, our results are not incompatible with the social capital explanation. How can future research advance this inquiry? First, to the extent that inequality in political participation is associated with lower provision of social goods,^{11,33} our hypothesis should be further tested by including additional policy variables (e.g., health care access, housing), with the expectation that doing so will “explain” the association we observed between unequal political participation and poorer self-rated health. Second, research is needed that uses measures of political inequality other than voting inequality (e.g., campaign financing). Last, testing the “social capital” mechanism will be more complex and will require the development of better measures of social capital. □

Contributors

T. A. Blakely led the designing of the study, the analysis and interpretation of the data, and the writing of the paper. B. P. Kennedy contributed to the design of the study, the interpretation of the data, and the writing of the paper. I. Kawachi conceived the study and contributed to the design, the interpretation of the data, and the writing of the paper.

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