Social Science & Medicine 69 (2009) 1333–1342

Contents lists available at ScienceDirect

Social Science & Medicine

journal homepage: www.elsevier.com/locate/socscimed

Rising U.S. income inequality, gender and individual self-rated health, 1972–2004☆

Hui Zheng*

Department of Sociology, Duke University, United States

ARTICLE INFO

Article history:
Available online 6 September 2009

Keywords:
USA
Income inequality
Self-rated health
Gender

ABSTRACT

The effect of income inequality on health has been a contested topic among social scientists. Most previous research is based on cross-sectional comparisons rather than temporal comparisons. Using data from the General Social Survey and the U.S. Census Bureau, this study examines how rising income inequality affects individual self-rated health in the U.S. from 1972 to 2004. Data are analyzed using hierarchical generalized linear models. The findings suggest a significant association between income inequality and individual self-rated health. The dramatic increase in income inequality from 1972 to 2004 increases the odds of worse self-rated health by 9.4 percent. These findings hold for three measures of income inequality: the Gini coefficient, the Atkinson Index, and the Theil entropy index. Results also suggest that overall income inequality and gender-specific income inequality harm men's, but not women's, self-rated health. These findings also hold for the three measures of income inequality. These findings suggest that inattention to gender composition may explain apparent discrepancies across previous studies.

Does income inequality harm health? This question has become increasingly important given the rise of income inequality in the U.S. over the last three decades. However, existing research provides divergent and uncertain evidence for a relationship between income inequality and health. In the 1990s, a wave of studies began to document the inverse relationship between income inequality and health (Kaplan, Pamuk, Lynch, Cohen, & Balfour, 1996; Kawachi, Kennedy, Lochner, & Prothrow-Stith, 1997; Subramanian, Delgado, Jadue, Vega, & Kawachi, 2003; Wilkinson, 1992). After that initial wave of research, however, a skeptical literature has begun to emerge that questions those initial findings (Beckfield, 2004; Judge, 1995; Lynch, Smith, & Harper, et al., 2004; Mellor & Milyo, 2001).

Although some of the proponents of the earlier studies responded to these skeptics and claimed these criticisms "would be both hasty and premature" (Subramanian, Blakely, Kawachi, 2003: 153), the field is now at a point of uncertainty. Indeed, it is perhaps not unreasonable to suggest that many scholars are skeptical of the effects of income inequality on health while others remain committed to the initial theoretical arguments and the possibility of such effects under certain empirical conditions.

Given this uncertainty, this paper will provide an alternative approach to examine whether income inequality harms health. Most current research is based on cross-sectional datasets and geographical variation (i.e., across nations or U.S. states). An alternative strategy is to see if temporal changes in income inequality affect health. The temporal comparison can provide a valuable alternative perspective that can more clearly determine whether there is a relationship from income inequality to health.

In this paper, I examine whether dramatic increases in income inequality in the U.S. have negatively affected individual self-rated health over the last three decades. I address this research question using both General Social Survey data and U.S. Census data from 1972 to 2004. Using Hierarchical Generalized Linear Models to capture both the individual and population levels, I examine whether income inequality has a contextual effect on individual health. More specifically, this research question is decomposed into three sub-questions: Do increases in income inequality have a negative effect on individual self-rated health? Do increases in income inequality predict gender-specific self-rated health? Do increases in gender-specific income inequality predict gender-specific self-rated health?

☆ I gratefully thank David Brady, Linda Burton, Linda K. George, Kenneth C. Land, and four anonymous reviewers for their very helpful comments and advice. In particular, both Linda George and David Brady have consistently made great contributions to the longtime revision of this paper. One reviewer's suggestion of the "natural experiment" analysis has significantly strengthened this paper.

* Tel.: +1 9194506620.
E-mail address: huizheng@soc.duke.edu

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doi:10.1016/j.socscimed.2009.08.016

Background

Income inequality and health

Wilkinson (1992, 1996) argues that in developed nations it is not the absolute standard of living that is important for health, but rather the depression, isolation, insecurity, and anxiety that are associated with income inequality. Subsequent studies documenting the inverse relationship between income inequality and health vary with regard to the health outcome examined and the unit of analysis. Some research examines how income inequality affects self-rated health (Blakely, Kennedy, Glass, & Kawachi, 2001; Lopez, 2004), and cardiovascular disease risk factors (Diez-Roux, Link, & Northridge, 2000) at the individual level. Other studies focus on the population level and examine whether income inequality affects life expectancy (Kawachi, Levine, Miller, Lasch, & Amick, 1994; Wilkinson 1992, 1996), and mortality rates (Kaplan et al., 1996; Kawachi & Kennedy, 1997).

Although it remains uncertain whether income inequality actually affects health, it is informative to note the potential pathways between them. According to previous research, the mechanisms linking income inequality to worse health include loss of social cohesion and the erosion of social capital (Kawachi et al., 1997; Wilkinson 1992, 1996), under-investment in human capital and other social goods (Davey, 1996; Kaplan et al., 1996), and the potentially harmful consequences of frustration brought about by relative deprivation (Brady, 2003; Kawachi et al., 1997; Stewart, 2006).

Explanations for mixed results

Although the negative relationship between income inequality and health has been observed in many studies, recent studies dispute previous findings, arguing that income inequality has no robust effect on population health (Deaton & Lubotsky, 2003; Judge, Mulligan, & Benzeval, 1998; Lynch et al., 2004; Mackenbach, 2003; Mellor & Milyo, 2001). For example, Beckfield (2004) criticized the cross-sectional analysis in previous research and did not observe an association between income inequality and population health (e.g., life expectancy and infant mortality rate) across 115 countries when using panel clustering of data. McLeod, Nonnemaker, and Call (2004) find no significant effects of income inequality on child well-being (e.g., infant mortality) in the 50 U.S. states when racial/ethnic composition is controlled.

Wilkinson and Pickett (2006) compiled a list of 168 analyses in 155 papers related to this topic and found that 87 analyses (52%) support the income inequality argument, 37 analyses (22%) did not support this argument and the rest, 26%, had mixed results. They also found that the amount of support for a relationship between income inequality and health differed by unit of analysis, declining from international comparisons, to large subnational areas and further to the smallest units (e.g., neighborhoods). They criticized small area studies because people's health is not affected by small area income inequality and relative deprivation within their neighborhoods, but rather by societal levels of income inequality and relative deprivation. After excluding small areas studies, 59% of the analyses support the income inequality argument, 23% do not, and 18% report mixed findings. I further examined these 168 analyses and identified two other patterns. First, studies based on U.S. samples consistently report stronger associations between income inequality and health than those based on samples from other countries. Second, the associations between income inequality and health do not significantly differ across specific health outcomes (e.g., mortality, life expectancy, diseases, self-rated health) (details available upon request). Therefore, use of self-rated health in this paper should not produce results that are widely discrepant from those that would be obtained for other health outcomes.

Although empirical analyses seem to slightly favor the income inequality argument, two other issues also substantially account for these divergent results: one issue is methodological, the other is theoretical. Some critics argue that previous studies of the impact of income inequality on population health have severe methodological shortcomings. First, previous studies relied primarily on bivariate or multivariate analyses without adequate statistical controls, which generated unobserved heterogeneity bias (Beckfield, 2004; Judge et al., 1998). The observed effects of income inequality may be due to variables that are confounded with income inequality, e.g., health expenditures, economic development for international comparisons and race composition for U.S. comparisons (Daly, Duncan, Kaplan, & Lynch, 1998; Mellor & Milyo, 2001). But some studies still find significant effect of income inequality on health after controlling for these possible confounding factors (e.g., Blakely, Kennedy, Glass, & Kawachi, 2001; Soobader & LeClere, 1999; Subramanian & Kawachi, 2004). Wilkinson and Pickett (2006) further claim that we need to differentiate genuine confounders from pathways or mediators and if some factors (e.g., race) are proxies for stratification, we should probably not control for these factors. Wilkinson and Pickett's argument is also related to sample size and statistical power: over-control of illegitimate variables may reduce the statistical power to detect the effect of income inequality on health. Second, some critics argue that cross-national comparisons may be inaccurate because of non-comparable data across countries (Beckfield, 2004; Judge, 1995). This problem can either produce an artificial association between income inequality and health or disguise the true relationship between them. Within-country studies may be a better approach to minimize the non-comparability problem involving cross-national studies.

In addition, aggregate studies of income inequality-population health suffer from theoretical ambiguity in terms of whether income inequality has compositional or contextual effects on health. Compositional effects refer to aggregate associations that result from accumulation of individual level associations, while contextual effects are aggregate associations resulting from the independent effect of social context (McLeod et al., 2004). Some critics argue that income inequality may have compositional effects on health and the aggregate association may result from the simple accumulation of individual level nonlinear associations between health and income (Gravelle, 1998), whereby "the redistribution of income from higher income groups to lower income groups (i.e., decreases in income inequality) will produce improvements in overall population health'' (McLeod et al., 2004: 251). Therefore, to determine whether income inequality has an independent contextual effect, potentially confounding individual level characteristics (e.g., income, education, age etc.) must be controlled. Micro-level data that match individual health status and other characteristics with macro-level information of income inequality are needed to test whether income inequality has contextual or compositional effects on individual health (Daly et al., 1998). Several studies have used multilevel data and methods and found that income inequality has contextual effects on individual health (e.g., Backlund et al., 2007; Torsheim, Candace, & William, et al., 2004).

An alternative approach

As noted above, focusing on a single country and using multilevel data may be the best approach for testing whether income inequality has contextual effects on health. Moreover, in this study I will combine this approach with a temporal approach, which has not been used in previous cross-sectional studies of geographic comparisons. Subramanian and Kawachi (2004: 87) suggested that "longitudinal observational data (e.g., repeated assessment of income inequality over time, in tandem with individual health outcomes) together with innovative application of multilevel structures may provide a better handle on the causal nature of the relation between income inequality and health." Temporal comparisons may be a valuable alternative perspective, because (1) it can provide a dynamic view of changes in income inequality vs. static comparison of levels of income inequality which, in turn, can better portray the causal relationship between income inequality and health; (2) compared to cross-national comparisons which cannot avoid many confounding factors (e.g., national characteristics), temporal comparisons can minimize the possible confounding factors although it cannot control all of them; (3) cross-national comparisons also have critical quality and comparability of data problems, due in part to different sampling procedures. By focusing on one country and using a temporal approach, these problems can be minimized. Therefore, this paper examines how temporal changes in income inequality in the U.S. over the past three decades affect individual health.

In addition, this paper also explores how income inequality affects gender-specific health outcomes. Some research suggests that relative deprivation strongly harms men's health through increasing stress and health compromising behaviors, and reducing self-esteem (Eibner & Evans, 2005), but is less harmful for women due to different gender roles in which men are embedded in more competitive settings and are the main breadwinners for families (Jones & Wildman 2008; Yngwe, Fritzell, & Lundberg, et al., 2003). This gender-specific relative deprivation-health pattern is not observed in Lorgelly and Lindley (2008). If relative deprivation harms men more than women, income inequality may harm men's health more than women's because relative deprivation is a potential mechanism linking income inequality to health (Kawachi et al., 1997). In addition, research finds that women report more and also are more vulnerable to interpersonal stressors or "network" events (i.e., life events that happen to significant others rather than the respondent) (Kessler & McLeod, 1984; Maciejewski, Prigerson, & Mazure, 2001). In contrast, men report more and are more vulnerable to legal and work-related stressful life events, e.g., job loss, legal problems, robbery, and work problems (Kendler, Thornton, & Prescott, 2001; Siegler & George, 1983). Income inequality is correlated with higher rates of violence, imprisonment, and unemployment (Kaplan et al., 1996), all of which involve more men than women, but may be less associated with interpersonal stressors, which disproportionably involve and are more harmful for women. Therefore, it is reasonable to hypothesize that income inequality will harm men's health more than women's. A recent study reports that, over the past three decades, men and women have different health trends (Hill & Needham, 2006). As shown in Fig. 1, self-rated health has improved steadily for women. But the pattern for men is less straightforward. Men's self-rated health declined during the 1970s, steadily improved during the 1980s, and then declined again during the 1990s. These findings suggest that continually rising income inequality from 1972 to 2004 may harm men's health more than women's—an issue unaddressed in previous research.

Methods

This analysis focuses on the U.S. population from 1972 to 2004, studying how changes in income inequality affect individual self-rated health, and requires a multilevel dataset. The individual level data for these analyses come from the National Opinion Research Center (NORC)'s General Social Survey (GSS), a pooled cross-sectional dataset. The macro-level data come from the *U.S. Census Bureau* (www.census.gov).

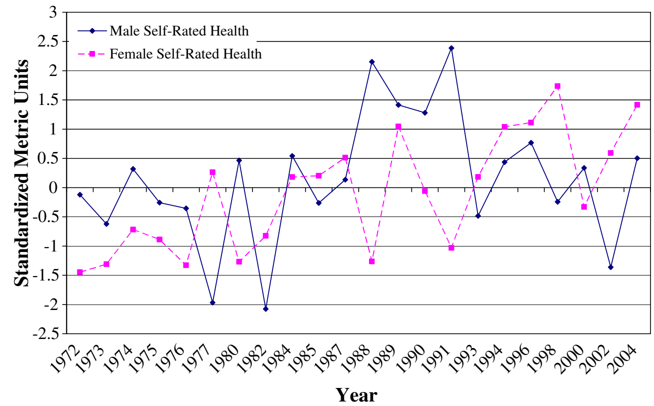

Fig. 1. Trends in Standardized Mean Self-Rated Health for Men and Women, 1972–2004. Note: Mean self-rated health has been standardized (0,1). Data Source: General Social Survey. Calculated by author.

The GSS involves face-to-face interviews. The GSS is a useful data source because of the consistency in question format over the years 1972–2004. Each year of the GSS is based on a national probability sample of noninstitutionalized US residents, 18 years of age and older. In total, the pooled GSS data have 25 waves. The dependent variable, self-rated health, has been included in most surveys since 1972 except 1978, 1983, and 1986. As a result, these analyses are based on 22 waves. I obtained data about income inequality from the U.S. Census Bureau. Three measures of income inequality are used: the Gini coefficient, the Atkinson index, and the Theil entropy index.

Measures

Outcome variable

Self-rated health. Respondents were asked: "Would you say your own health, in general, is excellent, good, fair, or poor?" On average, for the pooled sample across all waves, about 45% of respondents reported having "good" health, 32% "excellent", 18% "fair", and 5% "poor". The proportion of the population who describe their health as "good" or "excellent" has increased from 74.83% to 80.51% over the last three decades. A similar trend is observed for women, but the pattern for men is less straightforward, as reported in other research (Hill & Needham, 2006). Self-rated health is broader and more inclusive than more specific measures of health or impairment and extends beyond more objective measures of health. It is widely used to measure general health status and has been found to be very predictive of mortality and strongly correlated with objective assessments of health, including physician diagnoses (Idler & Benyamini, 1997). The relationships between self-rated health and objective health indicators also hold across population subgroups (Kennedy, Kasl, & Vaccarino, 2001). In the GSS, the response categories were coded excellent (1) to poor (4).[1]

Control variables

I also controlled for several variables that have been linked to health in previous research (e.g., Torsheim et al., 2004), including

[1] In Fig. 1, self-rated health was coded poor (1) to excellent (4). But in the following analyses, self-rated health was coded excellent (1) to poor (4) for ease of interpretation of income inequality's effect on health.

survey year, sex (1 = male, 0 = female), race (1 = white, others = 0), age in years, age squared, education, work status (1 = full/part time job, 0 = others), marital status (1 = married, 0 = others), and household income (see Appendix I). Survey year is recoded to range from 1 (1972) to 22 (2004). In the analyses, age squared is divided by 100 to reduce its scale. Education is measured as years of formal education completed. In the GSS, household income was measured by several income intervals. I first calculated the midpoint of each income interval and then adjusted these midpoint incomes for inflation based on the consumer price index (U.S. Bureau of Labor Statistics, 2006. The base period: 1982–1984 = 100). In the analyses, income is divided by 1000 to reduce its scale.

Explanatory variables

Income inequality. There are multiple ways to measure income inequality (Allison, 1978). Three measures of income inequality are used in this study: the Gini coefficient, the Atkinson index, and the Theil entropy index. All of them are calculated by the US Census Bureau.

The most commonly used measure of income inequality is the Gini coefficient, which is derived from the Lorenz curve, a curve that displays the cumulative distribution function of total income that accrues to successive income intervals. It ranges in value from 0 to 1, with higher values indicating higher levels of income inequality. In this study, overall income inequality is measured as the Gini coefficient of household income. Gender-specific income inequality is measured as the Gini coefficient of gender-specific individual income. From 1972 to 2004, the overall Gini coefficient increased from 0.401 to 0.466. From 1972 to 2000, the male-specific Gini coefficient and the female-specific Gini coefficient increased from 0.316 to 0.418 and from 0.271 to 0.345, respectively. Two other measures of income inequality are also used to test for robustness: the Atkinson index and the Theil entropy index. The Census Bureau provides these two indexes only for household income. Thus, gender-specific trends in inequality cannot be measured with these two indexes. From 1972 to 2000, the Atkinson index and Theil index of household income increased from 0.14 to 0.185 and from 0.279 to 0.404, respectively.

Statistical procedures

Because self-rated health is an ordinal valuable with four categories and the independent variables are located in two levels (individual level and population level), I used Hierarchical Generalized Linear Models (HGLM) to conduct the analyses. HGLM offers a coherent modeling framework for multilevel data with nonlinear structural models and non-normally distributed errors (Raudenbush & Bryk, 2002). Control variables from GSS are at the individual level and income inequality variables are at the population level. HGLM takes the following equations[2]:

Level-1 Model:

$$\log\left[\frac{P'}{(1-P')}\right] = \beta_0 + \beta_1(\text{SEX}) + \beta_2(\text{RACE}) + \beta_3(\text{AGE}) + \beta_4\left(\text{AGE}^2\right) + \beta_5(\text{EDUCATION}) + \beta_6(\text{MARITAL}) + \beta_7(\text{WORK}) + \beta_8(\text{INCOME}) + \varepsilon_0$$

Level-2 Model:

$$\beta_0 = \gamma_{00} + \gamma_{01}(\text{INCOME INEQUALITY}) + \varepsilon_1$$

[2] More detailed formulas are in the Appendix II.

The Level-1 model is basically an ordered logit model regressing self-rated health on several individual level control variables, where "excellent health" is the reference group. But the interpretation does not require specifying a given level of self-rated health which is compared to a specific reference group (Long, 1997). Instead, the interpretation can be put in this way: how the odds of reporting worse health are changed by a one unit increase in the explanatory variable. The Level-2 model is a linear regression model which assumes a normally distributed error term ε_1. It regresses the intercept β_0 in the Level-1 model on income inequality. The effects of the regressors (except the intercept β_0) in both models are treated as fixed rather than random effects. The intercept β_0 is a random coefficient which is associated with periods, such as the period-attributable variation in self-rated health ε_1. Income inequality is used to explain this period-attributable variation in health.

In the appendices, I display sensitivity analyses using ordered logistic model (OLM) with robust-clustered errors instead of HGLM. The results and conclusions are robust, and a replication of the analyses is displayed in the Appendices III–V for comparison.

The analysis proceeds in three steps. First, I examine the relationship between overall income inequality and individual health. Second, I examine the relationship between overall inequality and gender-specific health. The third step involves testing the relationship between gender-specific inequality and gender-specific health.

Results

Fig. 2 portrays the similar trends in the three measures of income inequality in the US from 1972 to 2004 after standardization. They have been basically continually rising in recent decades, which has been observed in a number of studies (e.g., Alderson & Nielsen, 2002). Moreover, the much sharper increase in income inequality from 1991 to 1993 and then persisting afterwards provides a "natural experiment" to examine the effect of income inequality on health, which will be addressed in more detail later.

Do increases in income inequality have a negative effect on individual self-rated health?

Panel A of Table 1 presents the odds ratios for HGLM from three models predicting self-rated health. Model 1 is the unconditional model without any predictors included. The second model controls for the level-1 individual explanatory variables (e.g., gender, race, age, SES). The third model additionally controls for level-2 population income inequality: the Gini coefficient of household income.

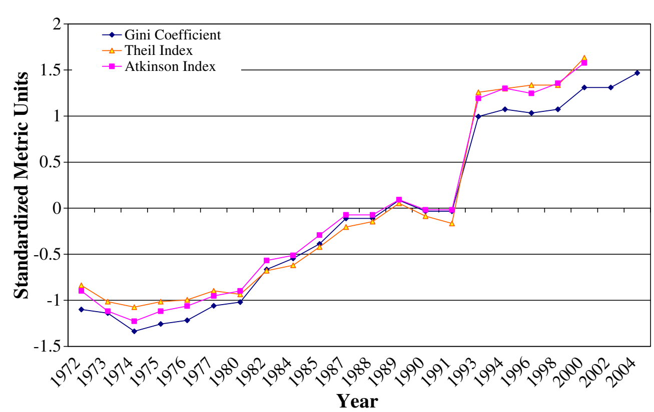

Fig. 2. Trends in Three Measures of Overall Income Inequality, 1972–2004: Gini Coefficient, Atkinson Index and Theil Entropy Index. Note: These three measures of income inequality have been standardized (0,1) to account for metric differences. Data Source: The Census Bureau.

Table 1
Odds Ratios for Hierarchical Generalized Linear Models of Self-rated Health on Income Inequality and Individual Predictors (Robust Standard Errors in Parentheses).

Panel A	Hierarchical Generalized Linear Models (Period 1972–2004)		
	Model 1 = Unconditional model	Model 2 = Level 1 Individual Predictors	Model 3 = 2 Levels/Gini
Fixed effects			
Intercept	.06*** (.05)	.08*** (.19)	.05*** (.11)
Sex (1 = man, 0 = woman)		1.06 (.03)	1.06(.03)
Race (1 = white, 0 = others)		.76*** (.03)	.76*** (.03)
Age		1.07*** (.01)	1.07*** (.01)
$Age^2/100$		.96*** (.01)	.96*** (.01)
Education		.89*** (.003)	.89*** (.003)
Marital (1 = married, 0 = others)		.90*** (.03)	.90*** (.03)
Work (1 = having job, 0 = others)		.62*** (.03)	.62*** (.03)
Income/1000		.99*** (.001)	.98*** (.001)
Income Inequality			
Gini			3.97** (.45)
Variance components			
Level 2: Across-Years	.0025**	.0017*	.0006
N	30819	30819	30819
Panel B	Ordered logistic model (Period 1991–2004)		
Gini			9.46* (1.09)

*** $p < .001$, ** $p < .01$, * $p < .05$.
Note: I use ordered logistic model with robust-clustered errors to estimate the effect of Gini on health instead of HGLM in Panel B due to small number of second level factor ($n = 8$). All individual level covariates in Panel A are also controlled in Panel B.

The results are very similar across models. Being white, highly educated, married, having a job and higher income reduce the odds of reporting worse health. The Gini coefficient predicts worse self-rated health. Every unit increase in the Gini coefficient substantially increases the odds of worse self-rated health by 2.97 (3.97 − 1) times. Thus, the dramatic increase in the Gini coefficient from 1972 to 2004 increased the expected odds of worse self-rated health by 9.4% ($(3.97^{0.065} - 1) \times 100$).[3] Although the effects of the Atkinson index and the Theil index are different from the Gini coefficient because of differences in metric, both predict worse self-rated health (i.e., the dramatic increase in the Atkinson index or the Theil index from 1972 to 2004 increased the odds of worse self-rated health by 10.4% and 3.9%, respectively) (table available upon request).

Another way to look at the effect of income inequality on health is to focus on how income inequality explains the period-variation in health. The variance components indicate the variation in self-rated health that is attributable to periods. Model 1 shows that the extent of unconditional period-attributable variation in self-rated health is about 2.5%. Individual sociodemographic markers reduce this residual variation to 1.7%, and so can account for about 28.6% (i.e., (2.5 − 1.7)/2.5) of the period-attributable variation. The Gini coefficient can account for about 64.7% (i.e., (1.7 − 0.6)/1.7) of the period-attributable variation in self-rated health. More important, after including the Gini coefficient, the variance components become non-significant, which means no further population level predictors are needed to explain the period-attributable variation in health. Thus, the Gini coefficient explains a great deal of the period variation in self-rated health.

Income inequality dramatically increased to a higher level between 1991 and 1993 and then persisted afterwards (see Fig. 2), which provides a "natural experiment" to examine if the significant rising income inequality in the period 1991–2004 has an even stronger impact on self-rated health than for the entire 22 years. If it is, then self-rated health is very sensitive to income inequality; if it is not, the effect of income inequality on health is less certain. The result in Panel B of Table 1 suggests a much stronger impact of income inequality on health between 1991 and 2004. The odds ratio is 9.46, much larger than 3.97 in HGLM (Panel A of Table 1) and 4.28 in Ordered Logistic Model (in Appendix III) for the whole period 1972–2004. This suggests that self-rated health is highly sensitive to changes in income inequality.

Fig. 3 shows the odds of worse self-rated health from 1972 to 2004 as predicted based on HGLM and standardized to the 1972 Gini coefficient and average socioeconomic levels (i.e., the control variables in Level-1 model). HGLM predicts that the odds of worse self-rated health decreased from 1972 to 2004 (as shown by the curve with circles). In other words, HGLM suggests a trend of better self-rated health over time. If the Gini coefficient had remained the same at 0.401 from 1972 to 2004, the odds of worse self-rated health would have declined faster than predicted by HGLM (as shown by the curve with squares). Simply put, increasing income inequality has suppressed the trend of increasing better self-rated health from 1972 to 2004. If socioeconomic characteristics had remained constant at the 1972 levels, the increasing Gini coefficient would have raised the odds of worse self-rated health over time (as shown by the curve with triangle dots). That means that if socio-economic development (e.g., education, income, and employment status) was not realized, the trend of increasing better self-rated health would be reversed to increasing worse self-rated health due to rising income inequality from 1972 to 2004.

Do increases in income inequality predict gender-specific self-rated health?

Panel A of Table 2 presents the odds ratios for HGLM from four models predicting gender-specific self-rated health. The results from models 1–2 show that for men, being white, highly educated, married, having a job and higher income reduce the odds of reporting worse health. Every unit increase in the Gini coefficient

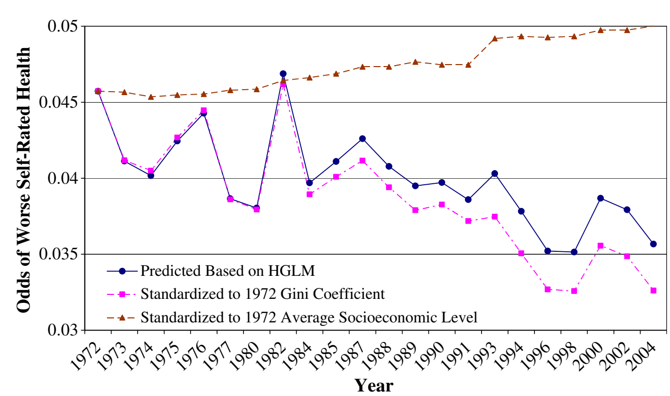

Fig. 3. Odds of Worse Self-Rated Health, 1972–2004 as Predicted and Standardized to the 1972 Gini Coefficient and Sociodemographic Levels. Note: I used a simple standardization technique by inputting the Gini coefficients and means of each control variable for each year inModel 3 in Table 1 to produce the odds ratios for each year. Standardized to 1972 Gini coefficient means keeping Gini constant at the 1972level while other variables change. Standardized to 1972 average socioeconomic level means keeping the means of individualsocioeconomic characteristics (e.g., education, income, work status, race, age, gender and marital status) constant at the 1972 level while Gini changes.

[3] As indicated in Table 1, every unit increase in the Gini coefficient increases the odds of worse self-rated health by 2.97 (3.97−1) times. The Gini coefficient is 0.466 in 2004, and 0.401 in 1972, so the change between 1972 and 2004 is 0.065. This 0. 065 unit increase in the Gini coefficient increases the odds of worse self-rated health by ($3.97^{0.065}$-1) (Long, 1997). This number multiplied by 100 is for ease of interpretation.

Table 2
Odds ratios for hierarchical generalized linear models of gender specific self-rated health on income inequality and individual predictors (robust standard errors in parentheses).

Panel A	Hierarchical generalized linear models (Period 1972–2004)			
	Men		Women	
	Model 1 = Level 1 Individual Predictors	Model 2 = 2 Levels/Gini	Model 3 = Level 1 Individual Predictors	Model 4 = 2 Levels/Gini
Intercept	.06*** (.16)	.02*** (.32)	.12*** (.15)	.12*** (.26)
Race (1 = white, 0 = others)	.80*** (.03)	.81*** (.05)	.74*** (.04)	.74*** (.04)
Age	1.08*** (.01)	1.08*** (.01)	1.06*** (.01)	1.06*** (.01)
Age^2/100	.95*** (.01)	.95*** (.01)	.97*** (.01)	.97*** (.01)
Education	.90*** (.01)	.89*** (.01)	.88*** (.01)	.88*** (.01)
Marital (1 = married, 0 = others)	.89** (.04)	.90** (.04)	.93* (.03)	.93* (.03)
Work (1 = having job, 0 = others)	.55*** (.05)	.55*** (.04)	.67*** (.03)	.67*** (.03)
Income/1000	.99*** (.000)	.98*** (.001)	.99*** (.000)	.98*** (.001)
Income Inequality				
Gini		17.40*** (.72)		1.03 (.73)
Variance components				
Level 2: Across-Years	.006**	.001	.003*	.003*
N	13986	13986	16833	16833
Panel B	Ordered Logistic Model (Period 1991–2004)			
Gini		1083.28*** (1.41)		.22 (2.59)

***$p < .001$, **$p < .01$, *$p < .05$.
Note: I use ordered logistic model with robust-clustered errors to estimate the effect of Gini on health instead of HGLM in Panel B due to small number of second level factor ($n = 8$). All individual level covariates in Panel A are also controlled in Panel B.

significantly increases the odds of men's worse self-rated health by 16.4 (17.4 − 1) times. The increase in the Gini coefficient from 1972 to 2004 increased the expected odds of men's worse self-rated health by 20.4% ($(17.4^{0.065} - 1) \times 100$). Both the Atkinson index and the Theil index have similar effects (table available upon request). The changes in variance components indicate that the Gini coefficient explains the period-attributable variation in men's self-rated health.

The results from models 3–4 show that for women, being white, highly educated, married, having a job and higher income reduce the odds of reporting worse health. The Gini coefficient, however, is not significantly related to self-rated health. Both the Atkinson index and the Theil index show similar patterns (table available upon request). The significant variance components also suggest that additional individual and population level predictors are needed to explain the period-attributable variation in women's self-rated health. In sum, Table 2 shows that overall income inequality worsens men's self-rated health but does not affect women's self-rated health. The bigger odds ratio of the Gini coefficient for men than for the overall population (model 2 in Table 2 vs. model 3 in Table 1) results from the elimination of women in Table 2. Because income inequality has no effect on women's health, including them in the analyses for the whole population reduces the odds ratio of income inequality on health.

Panel B of Table 2 shows the effect of the Gini coefficient on gender-specific health in the "natural experiment" period of 1991–2004. The Gini coefficient has a much stronger impact on men's health in this period, when the odds ratio is 1083, much larger than 17.4 in HGLM (Panel A of Table 2) and 17.62 in Ordered Logistic Model (in Appendix IV) for the whole period 1972–2004. In stark contrast, even in this period women's health is not affected by income inequality. Fig. 1 vividly portrays this pattern in this "natural experiment". Men's mean health steadily improved during 1987–1991 when income inequality was relatively stable, but then sharply declined from 1991 to 1993, coincident with dramatic increase in income inequality in these two years. After 1993, income inequality slowly increases and men's health slowly declines. Interestingly, GDP per capita consistently increased from $28,007 in 1991 to $28,940 in 1993 and $36,356 in 2004 (adjusted to 2000 dollars, source: *U.S. Bureau of Economic Analysis*), which definitely cannot explain the sharp decline of men's self-rated health in this period. But rising income inequality predicts men's health trend. In contrast, women's health continually increases in this period, which cannot be explained by changing income inequality.

Do increases in gender-specific income inequality predict gender-specific self-rated health?

I also examine whether gender specific self-rated health is affected by gender-specific income inequality. The results are very similar to Table 2, that is male income inequality harms men's self-rated health, but female income inequality is not associated with women's self-rated health (table available upon request). These analyses demonstrate that it is gender differences in the health outcomes rather than in income inequality that make a difference.

Discussion

Most previous research on income inequality and health is based on cross-sectional rather than temporal comparisons. Subramanian and Kawachi (2004) called for longitudinal studies of how income inequality affects individual health to more clearly portray the causal effect from income inequality to health. That approach is used in this study. Using General Social Survey and U.S. Census data, this study examined how changes in income inequality affect individual self-rated health in the U.S. from 1972 to 2004. By employing Hierarchical Generalized Linear Models that capture the effects of predictors operating at two levels (individual and population levels), a significant association between income inequality and individual self-rated health was observed even controlling for individual characteristics, which demonstrates income inequality's contextual rather than compositional effect on individual health, consistent with some previous research (e.g., Torsheim et al., 2004; Lopez, 2004). This study also finds that the increase in income inequality from 1972 to 2004 increased the odds of worse self-rated health by 9.4 percent. These findings hold for three measures of income inequality: the Gini coefficient, the Atkinson index, and the Theil entropy index. Income inequality dramatically increased to a higher level between 1991 and 1993 and then persisted afterwards. By employing this "natural experiment'' I find that the significant rising income inequality in the period 1991–2004 has an even stronger impact on self-rated health than in the whole period, which suggests self-rated health is highly sensitive to income inequality.

Another contribution of this study is the finding that the relationship between income inequality and health may actually be a reflection of gender-specific reactions to income inequality. Results show that income inequality harms men's, but not women's, self-rated health. This pattern is observed for both total income inequality and gender-specific income inequality. These findings shed light on a possible cause for the mixed findings in previous research – that is, the gender composition of the sample may affect the findings. For example, if the sample has a majority of women, a relationship between income inequality and health may not be found for the sample as whole. But these findings also raise the question of why the effect of income inequality on self-rated health is gender-specific.

Previous research suggests that three mechanisms link income inequality to worse health. First, income inequality causes loss of social cohesion and the erosion of social capital, which results in "unhealthy societies" (Wilkinson, 1992, 1996). But this impact may be more severe for men than for women. Men tended to have larger numbers and higher proportions of non-kin ties in their personal networks than women, even in similar social structural positions (Moore, 1990). It is possible that non-kin ties are more likely to be damaged by income inequality, while close family ties may not, which then contributes to the specific effect of income inequality on men but not women. Research suggests from 1985 to 2004 non-kin ties decreased more than kin-ties in the U.S. (McPherson, Smith-Lovin, & Brashears, 2006). Moreover, using the General Social Survey data, I find that women are significantly more satisfied with their friendships than men and this pattern becomes stronger from 1973 to 1994 (GSS dropped this variable after 1994) although it is only weakly significant (data not shown).

Second, income inequality causes under-investment in human capital and other social goods, because "in societies with rising inequalities, the interests of the rich begin to diverge from those of the typical family" (Kawachi & Kennedy, 1999: 221). The rich demand specialized services for themselves and are less likely to use and invest in public services (e.g., public education). The reduced social spending leads to reduction of life opportunities and deterioration of life circumstances for the public (Davey, 1996; Kaplan et al., 1996). But this pattern may be stronger for men than for women. Men are more likely to encounter and are also more vulnerable to work-related stressors while women are more likely to encounter and are more vulnerable to "network" stressors (i.e., life events that happen to significant others) (Kendler et al., 2001; Kessler & McLeod, 1984). It is possible that income inequality generates more work-related stressors than "network" stressors, which may make income inequality more harmful for men's than women's health.

Third, income inequality harms health by intensifying relative deprivation that induces psychological stress (e.g., negative self-assessment, frustration and depression), which can harm health (Kawachi & Kennedy, 1999). Compared to women, men are more embedded in competitive settings. A sense of relative deprivation should be stronger for men than for women. Rising income inequality can strengthen the sense of relative deprivation, making it more harmful for men's than women's health. My analysis with the GSS data suggests that unmarried employed women, who are probably in similar social settings to men, are harmed more by income inequality than other women although it is not statistically significant (data not shown). Overall, however, more research is needed to understand why income inequality is unrelated to women's health. We need to further test the possible mechanisms described above to see if they explain gender differences in the relationship between income inequality and health.

In this study, I did not control for macro economic development and other macro-level factors for both theoretical and practical reasons. First, I already control these factors (e.g., education, income, work status, race, age, gender and marital status) at the individual level. Aggregate level studies usually control for GDP per capita to account for economic development, and GDP per capita is an indicator for mean level of individual economic achievement. In this study, I control for individual level education, income and employment status, which not only account for the mean level of socioeconomic development (as demonstrated in Fig. 3 with the line representing "standardized to 1972 average socioeconomic level"), but more importantly account for the heterogeneity among individuals, which is the advantage of multilevel studies compared to aggregate level studies.

Second, theoretically, GDP per capita and other socioeconomic development should promote health while income inequality is expected to harm health. If income inequality carried the unobserved effect of macro socioeconomic development, it should either have a beneficial effect on health which is not true in this study, or be *underestimated* because macro socioeconomic development is positively related with the Gini coefficient and has a negative effect on worse self-rated health. As Fig. 3 suggests, socioeconomic development promotes health, but rising income inequality suppresses this trend. Indeed, rising income inequality wipes out all the gains from socioeconomic development on "excellent health" which leads to a fluctuating curve of "excellent health" around a constant level from 1972 to 2004 (figure available upon request).

Third, some studies show that states with larger proportions of African Americans have worse health outcomes; however, these states also have higher levels of income inequality (Deaton & Lubotsky, 2003). Therefore, the impact of racial composition may result from the damaging effect of income inequality on social welfare, public goods, human capital (Kaplan et al., 1996) and social cohesion (Kawachi et al., 1997) which cause worse health outcomes. Moreover, in this study, I examine temporal changes in income inequality for the whole U.S. rather than within states; it is not reasonable to assume that the whole country will reduce investments in social welfare because of over-representation of blacks or other minorities in some states. Therefore it is not necessary to control for racial composition in this study although it is already controlled at the individual level.

Fourth, empirically, in this study the Gini coefficient is highly correlated with GDP per capita, racial composition, and educational attainment at the macro level, $r = 0.971$, 0.973, 0.955, respectively. Including these factors in the analyses would cause severe degrees of freedom problems, collinearity problems and underpowered analyses, because the number of macro-level cases is just 20 or 22. Fifth, changes in variance components in Tables 1 and 2 suggest that no further population level predictors are needed to explain the period-attributable variation in health, at least for men.

This study suggests several directions for future research. First, this study is restricted to the U.S. Using a temporal approach to study the effect of changes in income inequality on health in other nations can further advance our understanding. Second, self-rated health is a widely used and reliable measure of general health status, but the effects of income inequality on other health outcomes should be examined. Third, this study examines the contextual effect of income inequality at the country level. It is not clear that country is the best level for understanding the contextual effects of income inequality. Further research about how the increases in income inequality from 1972 to 2004 for the 50 states in the U.S. affect individual health is encouraged, whereby a three-level model of individual, states, and time can be employed.

This study and others have demonstrated a negative impact of income inequality on individual self-rated health (e.g., Blakely et al., 2000; Kennedy, et al., 1998; Subramanian & Kawachi, 2003; Subramanian, Kawachi, & Kennedy, 2001). Although some previous research has claimed that this relationship is spurious by controlling for variables confounded with income inequality, it is not clear whether these variables are mediators of the effect of income inequality on health or if these studies are short of statistical power, causing Type II errors (i.e., the statistical model indicates no significant relationship although the relationship exists in reality). By neglecting this possible problem, however, our understanding of social processes may be compromised.

This study has demonstrated the negative contextual effect of income inequality on men's health. A "natural experiment" in 1991–2004 period suggests men's deteriorating self-rated health was largely caused by significantly rising income inequality while GDP per capita was still consistently increasing during this period. A policy implication of these findings is that reducing income inequality should improve men's self-rated health. Policy makers also need to recognize that improving the life conditions of the poor may not improve health if the life conditions of the rich are improved even more than those of the poor. Previous policies often target high risk individuals and groups (i.e., individuals and families living in poverty). This is a useful strategy for helping those at high risk, but we may be fighting a losing battle if we neglect the underlying societal forces (e.g., income inequality and the social costs of income inequality) that underlie poor health.

Appendix I. Descriptive statistics for dependent and key explanatory variables.

Outcome	Description	*N*	Mean	SD	Min	Max
HEALTH	Self-rated health. 1 = excellent, 2 = good, 3 = fine, 4 = poor	30819	1.965	0.842	1	4
Explanatory variables						
YEAR	Survey year	30819	1987	9.976	1972	2004
SEX	1 = man, 0 = woman	30819	0.454	0.498	0	1
RACE	1 = white, 0 = other races	30819	0.828	0.378	0	1
AGE	Respondent's age at survey year	30819	44.35	16.95	18	89
AGE^2	Age square	30819	2254	1676	324	7921
EDUCATION	Respondent's years of schooling	30819	12.62	3.161	0	20
MARITAL	1 = married, 0 = others	30819	0.566	0.496	0	1
WORK	1 = employed, 0 = others	30819	0.620	0.485	0	1
INCOME	Household income at survey year	30819	24930	16530	543	87873
Income Inequality						
GINI	Gini of household income	22	0.429	0.025	0.395	0.466
MALE GINI	Gini of male individual income	20	0.354	0.038	0.309	0.418
FEMALE GINI	Gini of female individual income	20	0.299	0.034	0.250	0.345
ATKINSON	Atkinson index of household income	20	0.156	0.018	0.134	0.185
THEIL	Theil entropy index of household income	20	0.321	0.051	0.267	0.404

Note: The Census Bureau provides gender-specific Gini coefficient, the Atkinson index and the Theil index only from 1972 to 2000.

Appendix II. : Formulas for HGLM models:

Level-1 Model:

$$\log\left[\frac{P'(1)}{(1-P'(1))}\right] = \beta_0 + \beta_1(\text{SEX}) + \beta_2(\text{RACE}) + \beta_3(\text{AGE}) + \beta_4\left(\text{AGE}^2\right) + \beta_5(\text{EDUCATION}) + \beta_6(\text{MARITAL}) + \beta_7(\text{WORK}) + \beta_8(\text{INCOME})$$

$$\log\left[\frac{P'(2)}{(1-P'(2))}\right] = \beta_0 + \beta_1(\text{SEX}) + \beta_2(\text{RACE}) + \beta_3(\text{AGE}) + \beta_4\left(\text{AGE}^2\right) + \beta_5(\text{EDUCATION}) + \beta_6(\text{MARITAL}) + \beta_7(\text{WORK}) + \beta_8(\text{INCOME}) + \delta_{(2)}(\text{THOLD2})$$

$$\log\left[\frac{P'(3)}{(1-P'(3))}\right] = \beta_0 + \beta_1(\text{SEX}) + \beta_2(\text{RACE}) + \beta_3(\text{AGE}) + \beta_4\left(\text{AGE}^2\right) + \beta_5(\text{EDUCATION}) + \beta_6(\text{MARITAL}) + \beta_7(\text{WORK}) + \beta_8(\text{INCOME}) + \delta_{(3)}(\text{THOLD3})$$

Level-2 Model:

$$\beta_0 = \gamma_{00} + \gamma_{01}(\text{INCOME INEQUALITY}) + \varepsilon_0$$

Combined Model:

$$\log\left[\frac{P'(1)}{(1-P'(1))}\right] = \gamma_{00} + \gamma_{01}(\text{INCOME INEQUALITY}) + \beta_1(\text{SEX}) + \beta_2(\text{RACE}) + \beta_3(\text{AGE}) + \beta_4\left(\text{AGE}^2\right) + \beta_5(\text{EDUCATION}) + \beta_6(\text{MARITAL}) + \beta_7(\text{WORK}) + \beta_8(\text{INCOME}) + \varepsilon_0$$

$$\log\left[\frac{P'(2)}{(1-P'(2))}\right] = \gamma_{00} + \gamma_{01}(\text{INCOME INEQUALITY}) + \beta_1(\text{SEX}) + \beta_2(\text{RACE}) + \beta_3(\text{AGE}) + \beta_4\left(\text{AGE}^2\right) + \beta_5(\text{EDUCATION}) + \beta_6(\text{MARITAL}) + \beta_7(\text{WORK}) + \beta_8(\text{INCOME}) + \delta_{(2)}(\text{THOLD2}) + \varepsilon_0$$

$$\log\left[\frac{P'(3)}{(1-P'(3))}\right] = \gamma_{00} + \gamma_{01}(\text{INCOME INEQUALITY}) + \beta_1(\text{SEX}) + \beta_2(\text{RACE}) + \beta_3(\text{AGE}) + \beta_4\left(\text{AGE}^2\right) + \beta_5(\text{EDUCATION}) + \beta_6(\text{MARITAL}) + \beta_7(\text{WORK}) + \beta_8(\text{INCOME}) + \delta_{(3)}(\text{THOLD3}) + \varepsilon_0$$

Appendix III. Odds ratios for ordered logistic regression of self-rated health on income inequality and other explanatory variables (robust standard errors in parentheses).

	Model 1	Model2	Model 3
Income inequality			
Gini	4.28** (.46)		
Atkinson		4.84* (.65)	
Theil			1.83** (.23)
Sex (1 = man, 0 = woman)	1.06 (.03)	1.04 (.03)	1.04 (.03)
Race (1 = white, 0 = others)	.76*** (.03)	.77*** (.03)	.77*** (.03)
Age	1.06*** (.01)	1.07*** (.01)	1.07*** (.01)
Age^2/100	.96*** (.01)	.96*** (.01)	.96*** (.01)
Education	.89*** (.004)	.89*** (.004)	.89*** (.004)
Marital (1 = married, 0 = others)	.90*** (.03)	.90*** (.03)	.90*** (.03)
Work (1 = having job, 0 = others)	.62*** (.03)	.62*** (.03)	.62*** (.03)
Income/1000	.98*** (.001)	.98*** (.001)	.98*** (.001)
Wald chi2	9245.4	11925	12519
N	30819	28035	28035

*** $p < .001$, ** $p < .01$, * $p < .05$.
Note: Since year is highly correlated with these three measures of income inequality, year was not put in the equation, but the analyses were clustered by years.

Appendix IV. Odds ratios for ordered logistic regression of gender specific self-rated health on income inequality and other explanatory variables (robust standard errors in parentheses).

	Male			Female		
	Model 1	Model2	Model 3	Model 4	Model5	Model 6
Income Inequality						
Gini	17.62*** (.71)			1.25 (.77)		
Atkinson		28.24*** (1.00)			1.02 (1.21)	
Theil			3.32*** (.352)			1.09 (.44)
Race (1 = white, 0 = others)	.80*** (.04)	.83*** (.03)	.83*** (.03)	.74*** (.05)	.72*** (.05)	.72*** (.05)
Age	1.08*** (.09)	1.08*** (.01)	1.08*** (.01)	1.06*** (.01)	1.06*** (.01)	1.06*** (.01)
Age^2/100	.95*** (.01)	.95*** (.01)	.95*** (.01)	.97*** (.01)	.97*** (.01)	.97*** (.01)
Education	.89*** (.01)	.89*** (.01)	.89*** (.01)	.88*** (.01)	.88*** (.01)	.88*** (.01)
Marital (1 = married, 0 = others)	.89** (.04)	.89** (.05)	.89* (.05)	.93* (.03)	.93* (.04)	.93* (.04)
Work (1 = having job, 0 = others)	.55*** (.06)	.56*** (.06)	.56*** (.06)	.67*** (.03)	.67*** (.03)	.67*** (.03)
Income/1000	.98*** (.001)	.98*** (.001)	.98*** (.001)	.98*** (.001)	.98*** (.001)	.98*** (.001)
Wald chi2	7816.1	8843.68	8754.52	4356.31	4164.38	4013.30
N	13986	12709	12709	16833	15326	15326

***$p < .001$, **$p < .01$, *$p < .05$.
Note: Since year is highly correlated with these three measures of income inequality, year was not put in the equation, but the analyses were clustered by years.

Appendix V. Odds ratios for ordered logistic regression of gender specific self-rated health on gender-specific income inequality and other explanatory variables (robust standard errors in parentheses).

	Male	Female
	Model 1	Model 2
Income inequality		
Gender-specific Gini	4.35** (.50)	.95 (.62)
Race (1 = white, 0 = others)	.83*** (.03)	.72*** (.05)
Age	1.08*** (.01)	1.06*** (.01)
Age^2/100	.95*** (.01)	.97*** (.01)
Education	.89*** (.01)	.88*** (.01)
Marital (1 = married, 0 = others)	.89** (.05)	.93* (.04)
Work (1 = having job, 0 = others)	.56*** (.06)	.67*** (.03)
Income/1000	.98*** (.001)	.98*** (.001)
Wald chi2	9124.90	4343.94
N	12709	15326

*** $p < .001$, ** $p < .01$, * $p < .05$.
Note: Since year is highly correlated with these three measures of income inequality, year was not put in the equation, but the analyses were clustered by years.

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