

Shifting dollars, saving lives: What might happen to mortality rates, and socio-economic inequalities in mortality rates, if income was redistributed?

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Abstract

Personal or household income predicts mortality risk, with each additional dollar of income conferring a slightly smaller decrease in the mortality risk. Regardless of whether levels of income inequality in a society impact on mortality rates over and above this individual-level association (i.e., the ‘income inequality hypothesis’), the current consensus is that narrowing income distributions will probably improve overall health status and reduce socio-economic inequalities in health. Our objective was to quantify this impact in a national population using 1.3 million 25–59-year-old respondents to the New Zealand 1996 census followed-up for mortality over 3 years.

We modelled 10–40% shifts of everyone’s income to the mean income (equivalent to 10–40% reductions in the Gini coefficient). The strength of the income–mortality association was modelled using rate ratios from Poisson regression of mortality on the logarithm of equivalised household income, adjusted for confounders of age, marital status, education, car access, and neighbourhood socio-economic deprivation. *Overall mortality* reduced by 4–13% following 10–40% shifts in everyone’s income, respectively. *Inequalities in mortality* reduced by 12–38% following 10–40% shifts in everyone’s income. *Sensitivity analyses* suggested that halving the strength of the income–mortality association (i.e., assuming our multivariable estimate still overestimated the causal income–mortality association) would result in 2–6% reductions in overall mortality and 6–19% reductions in inequalities in mortality in this New Zealand setting.

Many commentators have noted the non-linear association of income with mortality predicts that narrowing the income distribution will both reduce overall mortality rates and reduce inequalities in mortality. Quantifying such reductions can only be done with considerable uncertainty. Nevertheless, we tentatively suggest that the gains in overall mortality will be modest (although still potentially worthwhile from a policy perspective) and the reductions in inequalities in mortality will be more substantial.

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Introduction

There is strong international evidence for lower income being associated with poorer health status (Backlund, Sorlie, & Johnson, 1996; Blakely, Kawachi, Atkinson, & Fawcett, 2004; Bucher &

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Raglan, 1995; Ecob & Davey Smith, 1999; Lantz et al., 1998; Marmot, 2002; Martikainen, Makela, Koskinen, & Valkonen, 2001; McDonough, Duncan, Williams, & House, 1997; Sorlie, Backlund, & Keller, 1995). Furthermore, there is convincing evidence of a non-linear association of income with mortality such that each extra dollar of income buys a little less health gain (Backlund et al., 1996; Deaton, 2002; Ecob & Davey Smith, 1999; Gravelle, 1998; Subramanian & Kawachi, 2004; Wagstaff & van Doorslaer, 2000). Given this pattern it has been suggested that “raising the incomes of more disadvantaged people will improve the health of poor individuals” and “help reduce health inequalities” (Lynch et al., 2004). Moreover, as the health loss of the rich is expected to be less than the health gain of the poor for redistribution of incomes, the average health status of the population should increase (Gravelle, 1998). Our aim in this paper is to model changes in overall mortality rates and socio-economic inequalities in mortality that might arise from redistribution of income.

It is important to note that we are not addressing the ‘income inequality’ hypothesis, per se, in this paper. That is the hypothesis that a society with more equal income distributions will have better health outcomes for everyone, *over and above* that predicted by his or her personal (or household) income. This hypothesis posits positive ‘spill over’, contextual or ecologic effects, but is contested (Lynch et al., 2004; Subramanian & Kawachi, 2004). The strongest evidence is at the state-level in the United States, but there are non-confirmatory studies at regional levels in other countries—including New Zealand (Blakely, Atkinson, & O’Dea, 2003). In this current paper we address the health impacts of individuals moving up and down the income–mortality curve as predicted by the individual-level association of income with health, but we do not model a shift in the entire income–mortality association whereby a narrower income distribution confers an additional contextual advantage in lower mortality risks for all income groups.

It is also important to be cognisant at the outset of the many limitations of the modelling exercise presented in this paper. Researchers of the association of income with health (usually) base their interpretations on the implicit assumption that at least some of the observed association of income with health is causal, and—by extension—that changes in an individual’s income should result in

some change in health. The implicit recommendation for policy-makers is that income redistribution will reduce inequalities in health. However, when challenged as researchers to *quantify* the impact of income redistribution on overall population health and inequalities in health, we are not aware of any research that has provided such explicit estimates. We believe it is a legitimate role of researchers to at least estimate the likely health impacts of income redistribution. Indeed, other social epidemiologists are also taking tentative steps on quantitative health impact assessments of various income-related policy options (Cole et al., 2005). Whilst these estimates will inevitably be uncertain, and must come with an ‘uncertainty warning’, in our view the provision of such quantitative estimates sharpen the policy analysis and debate.

What are the key limitations of any modelling exercise of the impact of income redistribution on population health and inequalities in health? Many, although we will address five in particular: asking the right counterfactual or policy-relevant question; life-course determination of health; confounding of the observed income–health association that we base our modelling on; time lags between income change and health change; and the possible dead-weight costs to society of income redistribution.

Asking the right counterfactual or policy-relevant question: In this paper, we model the impacts of moving everyone’s income some ‘X%’ to the mean income. Policies that are to some extent redistributive are the norm in most developed countries, and setting a counterfactual question about different levels of such redistribution is not unrealistic: measures of income inequality vary between countries or over time within countries (Atkinson, 2003), and government policies directly or indirectly influence income distributions (e.g., taxation and welfare benefit policies). But is this the most likely policy action? Much, but not all, of the increase in income inequality in developed countries since the 1970s is the consequence of increased returns to education (Atkinson, 2003), and the state may not be readily able to undo or off-set these changes by way of taxation or welfare policies. Further, it may be more efficient to address income inequalities not by taxation and income redistribution per se, but by targeted provision of free welfare services such as education and health care. Nevertheless, an estimate of one possible policy mechanism—income redistribution—provides more information for policy-making and debate than hitherto existed.

Life-course determination of health: It is increasingly recognised that adult health is a function of a lifetime of exposures—indeed intergenerational histories (Kuh & Ben-Shlomo, 1997; Kuh, Ben-Shlomo, Lynch, Hallqvist, & Power, 2003). Regarding findings on income and health from longitudinal studies: “long-term income is more important for health than current income; income levels are more significant than income change; persistent poverty is more harmful for health than occasional episodes; and income reductions appear to have a greater effect on health than income increases” (Benzeval & Judge, 2001). Research on the US Panel Study of Income Dynamics, for example, has shown that persistently low income was particularly important for mortality risk, although income instability was also important (McDonough et al., 1997). We can perhaps distil the issues to two key questions: does a change in income cause a change in health, and by how much?; and what is the time lag between any change in income and a change in health status? We address these issues as confounding and time lags in the next two paragraphs.

Confounding: Confounding in epidemiology is defined as the mixing of effects whereby an exposure of interest has an association with an outcome of interest that is (in part at least) due to some correlated variable that predicts the outcome. It plagues observational studies. The gold standard methodology to estimate the unconfounded association of income with health, therefore, would be a randomised trial of income supplementation—however, such a study has not been conducted (Connor, Rodgers, & Priest, 1999). One alternative approach is to control for those variables that may be confounders of the income–health association in observational studies. For example, we have previously found that about half of the age and ethnicity-adjusted association of income with mortality was attributable to confounding factors (Blakely et al., 2004). But such analyses that control for confounders are still prone to error from either measurement error of the confounders or simply not including all potential confounders (Davey Smith & Phillips, 1990; Phillips & Davey Smith, 1991; von Elm & Egger, 2004). Another alternative approach is to use longitudinal studies with repeated measures on individuals that allow an assessment of how much a change in income predicts a change in health. We are aware of one such study that meets this latter requirement (McDonough & Berglund, 2003). Using the US

Panel Study of Income Dynamics, McDonough and Berglund estimated self-rated health as a function of persistent poverty, transient poverty and income to needs ratio. Whilst not a highlighted finding of their study, close inspection of their results in Table 6 demonstrates that the coefficient for changing income to needs (0.0101) was approximately 18% of the magnitude of the coefficient for income to needs at baseline (0.0559). That is, the impact of changing income to needs on contemporaneously changing self-rated health was about 20% of the magnitude of the baseline income estimate—controlling for transient and persistent poverty, education, race, marital status and age. Such a result is indicative only: the standard errors of the coefficients were approximately 10% of the coefficient magnitude, and no allowance has been made for time lags. But it does provide an indication that the fraction of the income–health association that is causal for *contemporaneously measured self-rated health and in the US setting* may be as low as 20%.

Time-lags: We are aware of no reliable quantitative studies of the time lag between income change and change in health status. However, for outcomes such as mortality there must be some elapsed time for an income change to alter one’s mortality risk, be it by stress, dietary or other pathways. Some of these pathways (as demonstrated in life course epidemiology) will take decades. However, not all causal mechanisms will take that long. For example, the rapid response of life expectancy to economic and social upheaval in the eastern European countries post 1989 point to the possibility for rapid health responses to socio-economic change (Men, Brennan, Boffetta, & Zaridze, 2003; Notzon et al., 1998).

Deadweight costs to society of income redistribution: Redistributing income is not a cost-free policy in that there are welfare reducing deadweight losses associated with tax-collection and re-distribution systems (Deaton, 2002). That is, whether due to tax avoidance or other inefficiencies, redistribution of income may lower total income or welfare to society. However, there are many complex issues involved (as detailed further in the Discussion section) and so we have used the simplifying assumption that there is no overall change in total income.

In this paper, we were able to use estimates of the income–mortality association adjusted for (measured) confounders, but we were not able to

quantitatively explore life-course determination, time lags, and dead weight costs.

Methods

This paper builds on work published elsewhere from the New Zealand Census-Mortality Study (NZCMS) (Blakely et al., 2004). Briefly, the shape and strength of the income–mortality association was estimated among four census-mortality cohorts formed by anonymous and probabilistic record linkage. In this paper, we just use the most recent census-mortality cohort (i.e., 1996–99). The age range was restricted to 25–59-year-olds, with the upper limit imposed to avoid problems with people retiring before age 65 with consequent drops in income. We excluded deaths and person-time in the first 6 months to overcome maximal health selection effects (i.e., where poor health prior to death causes a drop in income, thereby inducing reverse causation in the income–mortality association). We used total household income, equivalised for the number of children and adults in the household to allow for economies of scale. Individuals were then allocated to one of ten income categories (see Table 1), and Poisson regression¹ conducted to determine the rate ratios of mortality compared to the \$30,000–39,999 reference group. As ethnicity is a major determinant of both income and mortality in New Zealand, our baseline estimates of the ‘total’ income–mortality association were both age- and ethnicity-adjusted. Next, we made our best estimate of the unconfounded association of income with mortality, settling on a model that additionally adjusted for marital status, highest educational qualification, car access and small area socio-economic deprivation (Salmond, Crampton, & Sutton, 1998a,b). The latter measure of deprivation is based on census areas of approximately 100–150 people, and uses data on the proportion of people meeting certain characteristics according to means-tested benefits, household income, unemployment, telephone access, car access, qualifications, tenancy, household crowding and single parent homes. It is calculated in much the same way as the Carstairs or Townsend

indices (Carstairs, 1995). Whilst labour force status is a major determinant of income, and a determinant of mortality, adjusting for labour force status was problematic, probably due to labour force status also being a proxy for health status — a variable obviously between income and mortality on any causal pathway.

Finally, we assessed the shape of the income–mortality association. It was clearly non-linear with income expressed simply as dollars. Both a rank and logarithm transformation of income improved the fit of the income–mortality association. Due to a probably aberrant dip in mortality rates for very low incomes (e.g., due to self-employed people with no declared income), the log transformation did not appear to fit particularly well at low incomes. However, we believe the log transformation is a good option for modelling the effects of income redistribution, and we exploit it in this paper just as others have done previously (e.g., Wolfson, Kaplan, Lynch, Ross, & Backlund, 1999).

To estimate the impact of income redistribution on both overall mortality rates, and relative inequalities in mortality rates, we specified two types of counterfactuals or hypothetical interventions. First, we modelled lifting everyone living on an equivalised household income of less than \$20,000 (14% of males, 20% of females) to have an equivalised household income of \$20,000–24,999 (mean income \$22,540). The ‘cost’ of this growth in total household income was 3–4% of the total male and female household income, respectively, or equivalent to New Zealand’s current rate of annual economic growth (www.stats.govt.nz). Second, we modelled various shifts in the entire income distribution, where everybody’s household income moved a specified percentage towards the mean household income. This particular scenario has two notable features: the Gini coefficient (a common measure of income inequality) reduces by the same percentage (see Appendix for proof), and the total of all household incomes remains the same (i.e., we are purely redistributing income). For both these counterfactual scenarios, we utilised the coefficient for the logarithm of income (specified as a continuous variable) in the multivariable Poisson model specified above.

We used the populations attributable risk percent (PAR%) (Murray & Lopez, 1999) to estimate the impact on total population mortality. This PAR% calculation entailed calculating the new relative risks of mortality for each of the ten income groups

¹Poisson regression is the standard epidemiological and biostatistical practice for regression modelling of a dependent variable involving count data (i.e. number of deaths) with a time-related denominator (i.e. person-years of observation within which the counts were observed). Poisson regression is generalised linear model, using a log-link in the same manner as logistic or probit regression – but with a Poisson distributed error-term.

Table 1

Person-years, deaths and rate ratios of mortality by equivalised household income among 25–59-year-old deaths 1996–99

Equivalised household income	Person-years	Weighted deaths	Observed rate ratios, using categorical income data (95% CI)		Predicted rate ratios, using coefficients for logarithm of income	
			Age- and ethnicity-adjusted	Multivariable ^a	Age- and ethnicity-adjusted	Multivariable ^a
<i>Males</i>						
≥ \$70,000	278,772	504	0.64 (0.56–0.73)	0.76 (0.67–0.88)	0.59	0.74
\$60,000–69,999	149,472	267	0.65 (0.56–0.74)	0.72 (0.63–0.83)	0.74	0.84
\$50,000–59,999	191,736	441	0.81 (0.71–0.94)	0.87 (0.76–1.01)	0.81	0.88
\$40,000–49,999	245,037	585	0.85 (0.75–0.97)	0.89 (0.78–1.02)	0.89	0.93
\$30,000–39,999	267,930	747	1	1	1.00	1.00
\$25,000–29,999	153,042	468	1.12 (0.97–1.28)	1.06 (0.92–1.22)	1.13	1.07
\$20,000–24,999	112,731	321	1.09 (0.93–1.27)	1.03 (0.88–1.21)	1.24	1.14
\$15,000–19,999	111,408	468	1.43 (1.25–1.64)	1.23 (1.07–1.42)	1.39	1.21
\$10,000–14,999	80,334	384	1.57 (1.35–1.82)	1.24 (1.06–1.44)	1.63	1.33
< \$10,000	44,205	183	1.40 (1.15–1.70)	1.21 (0.99–1.47)	2.92	1.88
Total	1,634,667	4368				
<i>Females</i>						
≥ \$70,000	258,813	252	0.60 (0.50–0.71)	0.69 (0.58–0.83)	0.61	0.73
\$60,000–69,999	140,817	171	0.74 (0.62–0.89)	0.81 (0.68–0.97)	0.75	0.83
\$50,000–59,999	183,177	282	0.91 (0.77–1.08)	0.97 (0.82–1.15)	0.82	0.88
\$40,000–49,999	238,116	348	0.89 (0.76–1.04)	0.92 (0.79–1.08)	0.89	0.93
\$30,000–39,999	272,442	447	1	1	1.00	1.00
\$25,000–29,999	159,678	321	1.19 (1.01–1.39)	1.14 (0.97–1.34)	1.12	1.08
\$20,000–24,999	126,954	246	1.29 (1.08–1.54)	1.22 (1.02–1.45)	1.22	1.14
\$15,000–19,999	144,975	366	1.43 (1.22–1.67)	1.27 (1.08–1.49)	1.37	1.22
\$10,000–14,999	125,268	324	1.66 (1.41–1.95)	1.38 (1.16–1.63)	1.59	1.35
< \$10,000	75,183	135	1.27 (1.02–1.58)	1.07 (0.86–1.35)	2.40	1.75
Total	1,725,423	2892				

Raw numbers are random rounded to the nearest multiple of three as per the Statistics New Zealand protocol.

^aIn addition to age and ethnicity, the multivariable model includes marital status, highest educational qualification, car access and small area socio-economic deprivation. *Source:* Blakely et al., 2004.

(as grouped *before* income redistribution) using the newly estimated counterfactual mean income for each group.

We also estimated the change to relative inequalities in mortality for each counterfactual. To avoid findings based on extreme income groups, we used the rate ratio for the second to lowest income group (\$10,000–14,999) compared to the second to highest income group (\$60,000–69,999) based on the actual groupings prior to any counterfactual. (Noting the distribution of person-years in Table 1, this equates to the relative risk of mortality for, approximately, the 95th percentile of household income compared to the 20th percentile.) We used the pre-counterfactual rate ratio, adjusted for age and ethnicity, comparing these two groups as our baseline measure of relative inequality in mortality. The change in the multivariable rate ratio for each category due to changing incomes was then used to estimate what the age and ethnicity-adjusted rate

ratio might be in the counterfactual. For example, if the baseline age and ethnicity-adjusted ratio was 2.0, and the multivariable rate ratios were 1.6 and 1.5 before and after income redistribution, then our estimated age and ethnicity-adjusted rate ratios for the counterfactual scenario was $2.0 - (1.6 - 1.5) = 1.9$. Finally, the percentage change from 2.0 to 1.9 was calculated using excess rate ratios (i.e., the rate ratio minus 1.0) given that the null is a rate ratio of 1.0. Continuing this example of a reduction in the rate ratio from 2.0 to 1.9, therefore, the percentage reduction was 10%.

Results

Table 1 shows the weighted person-years and deaths, and observed rate ratios for each of the ten categories of household income in the 1996–99 cohort. The coefficients for the logarithm of household income from Poisson regression models that

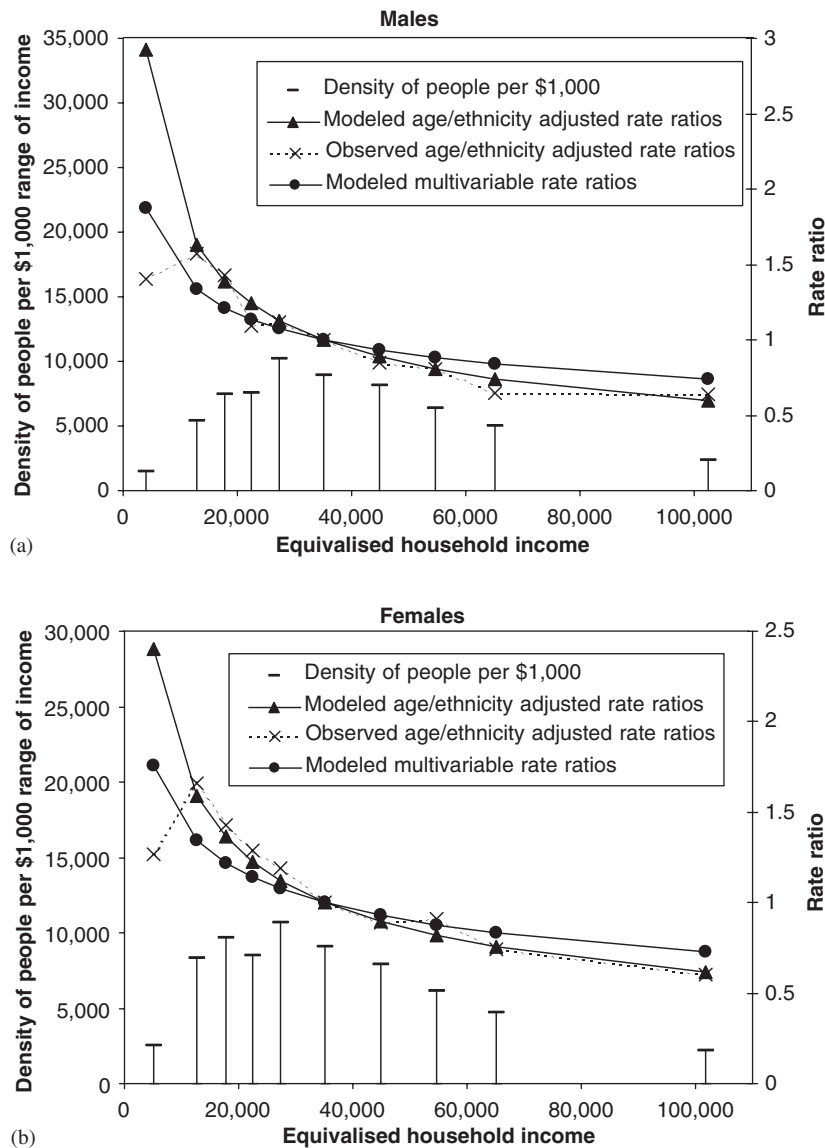


Fig 1. Density of people per \$1000 range of household income, and rate ratios of mortality (reference group \$30–39,999) for various specifications of the household income variable, 25–59-year-olds during 1996–99.

adjusted for age and ethnicity were -0.488 and -0.457 for males and females, respectively. The coefficients from the multivariable models were -0.286 and -0.293 . (These coefficients were reasonably similar for the earlier cohorts in the NZCMS (1981–84, 1986–89, 1991–94), using inflation-adjusted income data. Results available from authors on request.) Using the mean income in each income category, the predicted rate ratios by income category are also shown in Table 1.

Table 1 shows the density of census respondents per \$1000 increment in household income (left hand

y-axis). For both males and females, the classic right skewed income distribution is evident. Superimposed on Fig. 1 are some of the rate ratios presented in Table 1. Except for the lowest income category, there is a good visual fit of the log-transformed income association to the observed categorical rate ratios. The predicted rate ratios from the log-transformed income variable in the multivariable model are also plotted; the attenuation due to confounding is clear.

Table 2 presents results for the counterfactual scenarios. Focusing first on the estimations from the

Table 2

Percentage reductions in overall mortality rates (population attributable risk percent (PAR%)) and changes in the relative risk of mortality by income for various income redistributions

Change of income distribution modelled	Resulting % reduction in Gini coefficient	Using coefficient of the logarithm of household income from the preferred multivariable model ^a		Sensitivity analysis: Halving the coefficient of the logarithm of household income from the preferred multivariable model ^a	
		% Reduction in total mortality rate (ie, PAR%)	Estimated age-& ethnicity-adjusted rate ratio for people in the 2nd lowest c.f. 2nd highest income group (% reduction of relative inequalities) ^b	% Reduction in total mortality rate (ie, PAR%)	Estimated age- & ethnicity-adjusted rate ratio for people in the 2nd lowest c.f. 2nd highest income group (%reduction of relative inequalities) ^b
Males					
Do nothing	0	0	2.21 (0)	0	2.21 (0)
A: People with income <\$20,000 raised to \$20–\$24,999 category	9.9	3.5	1.94 (22)	1.5	2.09 (10)
B: Everyone's household income moves 'X' % to the mean					
household income					
X = 10%	10	3.7	2.06 (12)	1.7	2.14 (6)
X = 20% ^c	20	6.6	1.95 (22)	3.1	2.08 (10)
X = 30% ^d	30	9.2	1.85 (29)	4.4	2.03 (15)
X = 40%	40	11.7	1.77 (36)	5.6	1.99 (18)
Females					
Do nothing	0	0	2.11 (0)	0	2.11 (0)
A: People with income <\$20,000 raised to \$20–\$24,999 category	13.9	4.8	1.83 (25)	2.1	1.99 (11)
B: Everyone's household income moves 'X' % to the mean					
household income					
X = 10%	10	3.7	1.97 (13)	1.8	2.04 (6)
X = 20% ^c	20	7.3	1.86 (22)	3.4	1.99 (11)
X = 30% ^d	30	10.2	1.77 (31)	4.9	1.94 (15)
X = 40%	40	12.9	1.69 (38)	6.3	1.90 (19)

^aThe preferred multivariable model adjusted for age, ethnicity, marital status, highest educational qualification, car access and small area socio-economic deprivation. (See Methods for further details.)

^bThe estimated age- and ethnicity-adjusted rate ratio was calculated as described in methods. The percentage reduction was that for the excess rate ratio (i.e., RR minus 1), using the “Do nothing” or baseline age and ethnicity adjusted rate ratio as the baseline rate ratio.

^cBetween the mid-1980s and mid-1990s, the Gini in New Zealand increased by approximately 20% - although exact estimates vary by data-sets and analyses (Förster & d'Ercole, 2005; O'Dea, 2000; Statistics New Zealand, 1999). Therefore, the 20% reduction in Gini scenario has an equivalence to returning to 1986 Gini levels.

^dA 30% reduction in New Zealand's 1996 Gini would lead to a Gini similar to that for Sweden at the same time.

preferred multivariable model, overall mortality rates might decrease by 4–13% for 10–40% shifts in everyone's household income to the mean, respectively. But relative inequalities, as measured by the excess relative risk for the second to lowest (approximately 95th percentile) to second to highest (approximately 20th percentile) income groups, might decrease by 12–38% for 10–40% shifts in income, respectively. The targeted approach of lifting people out of the lowest incomes to the next highest income category reduces total mortality by around 4%, but reduces relative inequalities by 22% (males) to 25% (females).

Also shown in Table 2 are sensitivity analyses where we halved the coefficient for log transformation of income. That is, we assumed that our best multivariable model still overestimated the strength of causal association of income with mortality by two-fold (e.g., due to failure to adjust for labour force status). Essentially, all estimates (PAR% and reductions in relative inequalities) were approximately halved. Should one wish to be more conservative still, say setting the 'true causal' income mortality association at 20% of our 'preferred model', then the percentage estimates (both PAR% and reduction in relative inequalities) correspondingly reduced to about 20% of those from our preferred model.

Discussion

An important role of social epidemiology is to inform policy debates on reducing inequalities in mortality with, where possible, quantified effects. Many researchers have pointed to the non-linear association of income with mortality as a win-win scenario—narrowing income distributions will both improve overall mortality, and reduce inequalities (Gravelle, 1998; Kawachi, 2000; Lynch et al., 2004; Subramanian & Kawachi, 2004). Our modelling supports this argument, but makes it clear that the percentage gains in overall mortality will be less than the percentage reductions in inequalities in mortality. Whilst the income–mortality association is non-linear, redistributing income away from those with above average incomes still has some small detrimental health impact. Nevertheless, a modest reduction in overall mortality rates is still a major improvement in health status (e.g., a 5% reduction in all cause mortality among 25–59-year-olds would approximate preventing all unintentional injury deaths).

The reductions in relative inequalities in mortality, however, are sizeable. A 30% reduction in the Gini coefficient (the same relative distance between New Zealand's and Sweden's Gini coefficient (Statistics New Zealand, 1999)) might reduce relative inequalities by up to 30%. A return to an early 1980s pattern of income distribution in New Zealand (an approximately 20% reduction in the 1996 Gini (Förster & d'Ercole, 2005; O'Dea, 2000; Statistics New Zealand, 1999)) might reduce relative inequalities in mortality by up to 9%. And a targeted approach of lifting the incomes of the three bottom income groups to that of the fourth to bottom income group ('cost' equivalent to 3–4% of total household income) might reduce relative inequalities in mortality by up to a quarter (Table 2).

As indicated in the Introduction, there are numerous limitations to our modelling. First, we have produced our best relative risk estimate of the causal association of income with mortality based on adjusting for likely confounders, using perhaps one of the best national data-sets available for such estimations. However, 'declaring independence', or in this case the 'causal rate ratio', is a difficult endeavour in observational epidemiology (Davey Smith & Phillips, 1990; Phillips & Davey Smith, 1991; von Elm & Egger, 2004). There are likely to be unmeasured or unobservable risk factors that leave residual confounding (Deaton, 2002). Relevant to this analysis is labour force status, without doubt a determinant of income and probably a determinant of mortality risk independent of income and the other covariates in our preferred multivariable model. However, including labour force status in the model probably over-adjusts due to labour force status being a proxy for health status (Blakely et al., 2004). On the other hand, the inclusion of car access or small area socio-economic deprivation in our model might be considered over-adjustment since one's income influences both car ownership and where you live, putting these variables (in part at least) on the causal pathway. Nevertheless, even if we are conservative and specify the strength of causal income–mortality association as half what we estimated, there were still substantial reductions in relative inequalities in mortality (Table 2). (If we are even more conservative, and specify the strength of the causal association as only 20% of our best multivariable estimate (i.e., to be consistent with the McDonough & Berglund (2003) paper described in the Introduction), the PAR% and percentage

reduction in inequality estimates reduce further in a proportional manner).

As indicated in the Introduction, a life-course model of disease causation and time lags suggest any improvements in overall mortality, and reductions in inequalities, will take years and probably decades to fully accrue (Murray & Lopez, 1999). We simply do not have enough information to estimate how quickly gains will be made after income redistribution.

Finally, redistributing income may not be a cost-free policy due to economic dead-weight costs associated with some forms of tax-collection and re-distribution (Deaton, 2002)—we have not included this in our modelling. In addition to the difficulty to model deadweight costs, it is also not clear what magnitude these deadweight costs would be. For example, there is potential for a greater reliance in the New Zealand taxation system on Pigovian taxes (i.e., taxes imposed because of negative externalities for, say, alcohol). Any dead-weight cost of certain excise taxes may be outweighed by their benefits in reducing the adverse externalities (e.g., health benefit from reduced air pollution, reduced private car use, reduced tobacco and alcohol consumption). Another issue is that spatial variation in income redistribution policies is not modelled here and a national perspective for a small country is taken (in contrast to larger countries with both federal and state systems such as the USA).

What can we conclude from this study? Whilst our estimates obviously come with considerable uncertainty, they are perhaps the best that we can quantify with currently available data. The actual extent of changes in overall mortality, and inequalities in mortality, will also depend on contextual factors such as national health and welfare policies. Life-course determination of disease and time lags will inevitably mean any (full) return is years down the track from any change in income distribution. On the other hand, if there are positive spill-over effects on the health of the population due to lower levels of income inequality (i.e., the income inequality hypothesis is true), then the health gains will be greater than suggested in this paper that just models changes at the individual level. Nevertheless, the estimates in this paper tell us that, yes, income redistribution will probably result in a modest reduction in overall mortality rates — but not a large one, especially if the dead-weight costs of redistributing income are significant. They also tell

us that income redistribution should reduce relative inequalities in mortality.

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Appendix

The Gini coefficient is given by the area between the Lorenz curve and a straight line diagonal on a graph of cumulative income by people ranked from lowest to highest income, divided by the area beneath the diagonal. That is, in Fig. 2 the Gini coefficient is given by the area between the dashed diagonal and the curve, divided by the triangular area beneath the dashed diagonal.

For a large number of people in a population, the area between the diagonal and the Lorenz curve approximates:

$$\begin{aligned} \text{Area} &= \sum^i \left(i\bar{x} - \sum^i x_i \right) \\ &= \bar{x} \sum^i i - \sum^i \sum^i x_i \end{aligned}$$

where there are i individuals in the population, ranked in ascending order of income, the i th person's income is given by x_i and \bar{x} is the mean income.

Put in words, the first equation above sums across all individuals the short fall between the diagonal line where everyone receives the same (i.e., mean) income (i.e., $i\bar{x}$), and the actual cumulative income to that point (i.e., $\sum^i x_i$).

Under a counterfactual scenario where every person's income moves a constant proportion to the mean income, the counterfactual income for the i th person is

$$x'_i = (x_i + z(\bar{x} - x_i)),$$

where z is the proportion movement of everyone's income to the mean.

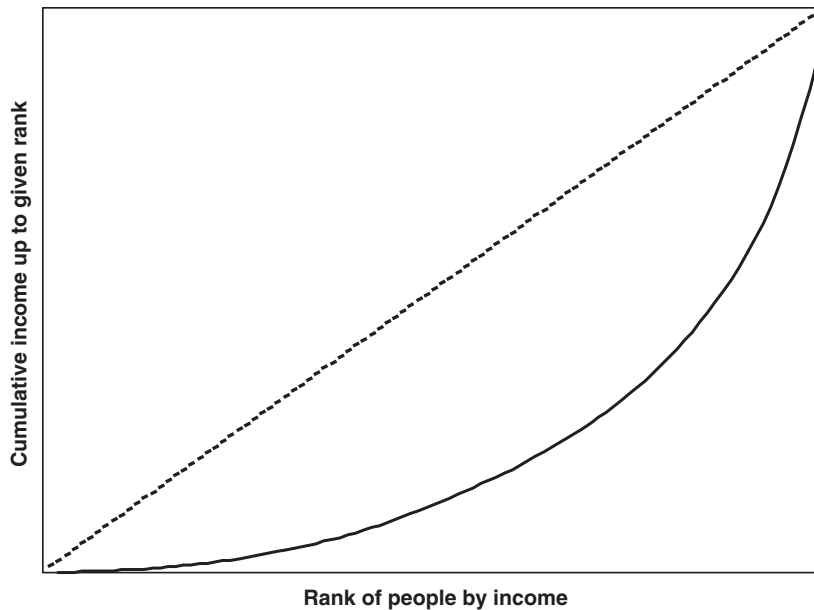


Fig 2. Lorenz curve.

Then the new area is

$$\begin{aligned}
 \text{Area}' &= \sum^i \left(i\bar{x} - \sum^i (x_i + z(\bar{x} - x_i)) \right) \\
 &= \bar{x} \sum^i i - \sum^i \sum^i (x_i + z(\bar{x} - x_i)) \\
 &= \bar{x} \sum^i i - \sum^i \sum^i x_i - z \left(\bar{x} \sum^i i - \sum^i \sum^i x_i \right) \\
 &= (1 - z) \left(\bar{x} \sum^i i - \sum^i \sum^i x_i \right) \\
 &= (1 - z) \text{Area}.
 \end{aligned}$$

That is, a percentage shift of $(z \times 100\%)$ in everyone's income to the mean income also reduces the Gini coefficient by the same percentage amount.

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