Changes in the Value of Life, 1940-1980

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ABSTRACT

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We present the first nationwide value of life estimates for the United States at more than one point in time. Our estimates are for every ten years between 1940 and 1980, a period when declines in fatal accident rates were historically unprecedented. Our estimated elasticity of value of life with respect to per capita GNP is 1.5 to 1.7. We illustrate the importance of rising value of life for policy evaluation by examining the benefits of improved longevity since 1900. Our estimated elasticity implies that the current marginal increase in longevity is more valuable than the large increase in the first half of the twentieth century.

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1 Introduction

Job safety improved sharply after World War II. Between 1940 and 1980 fatal workplace accidents declined by 0.08 per million hours worked, with roughly 90 percent of the decline occurring prior to 1960. By historical standards this decline was unprecedented. Fishback (1992: 103) reports that in coal mining fatality rates were constant between 1904 and 1930. The 1930s experienced a slight downward trend in fatality rates in the mineral industries, but nothing as pronounced as the 1940-1980 decline that we document (*Minerals Yearbook*). Although railroad fatality rates fell during the 1910s and 1920s, they were roughly constant between 1894 and 1911 and during the 1930s and 1940s (Fishback and Kim 1993). But, even though the largest changes in fatality rates occurred in the first two decades after World War II, most researchers measuring the value of life have generally focused on the 1970s onwards (see Viscusi and Aldy (2003) and Viscusi (1993) for a review).

This paper presents the first nationwide value of life estimates for the United States at more than one point in time. We estimate value of life in 1940, 1950, 1960, 1970, and 1980 using census micro-data and BLS fatality data by industry to recover the trend in the value of life. These years provide us with consistent data series and, more importantly, shed light on a little studied but important period for job safety.

Our hedonic regressions estimated at five points in time yield information on how compensation for job risk has changed as job safety has risen. We document that as the quantity of safety has increased over time, the compensating differential has also increased. This is strong evidence that the demand for safety has increased.

Repeat hedonic regressions are also useful for establishing the incidence of safety improvements over time. We document which socioeconomic and demographic groups have had the greatest reduction in risk exposure over time and we decompose reductions in average job risk into composition effects versus within industry risk progress.

Our findings have implications for the use of value of life estimates by government agencies for prospective policy evaluation and by academics for retrospective policy evaluation. A rising value of life suggests that marginal improvements in safety and in longevity are becoming more valuable. Most analysts and researchers, however, use value of life estimates derived from 1970s and 1980s data on compensating wage differentials for job risk, commonly measuring the value of life as between 3 to 6 million 1990 dollars. Although government agencies, such as the Environmental Protection Agency, have recently begun to make income adjustments for the value of life, researchers who have estimated the benefits of increased longevity and health over several decades or even over a whole century have treated the value of life as a constant (e.g. Nordhaus 2002; Murphy and Topel 2002; Cutler and Richardson 1997).

2 Empirical Framework

There is no reason to think that willingness to pay for risk reduction has remained constant when incomes have risen and the risk of death from other causes has fallen. On theoretical grounds we would expect that the value of life has increased because incomes have increased. Rosen's (1988) model shows that the value of life can be expressed as the marginal rate of substitution between wealth and the probability of survival. Unless people pay to increase risk, as wealth increases so does the value of life. For a working person, as the wage increases so does the value of life.

The value of life will also increase with improvements in elderly health, longevity, and well-being.[1] Suppose that individuals live for two periods, working when young and then, if they

[1] Dow, Philipson, and Sala-i-Martin (1999) show that a decline in the probability of death from other causes will increase value of life.

survive, retiring. An individual will maximize

$$U(w(r)) + (1-r)(1-s)V\,, \tag{1}$$

where r is the probability of dying on the job, $w(r)$ is the wage give risk r, U is utility, s is the probability of dying from disease, and V is discounted utility from retirement. In the United States the one year mortality rate at age 65 fell from 0.04 in 1900 to 0.03 in 1940 and then to 0.02 in 1980. As s falls, the risk premium, $w'(r)$, will increase as the young value living more. If V increases because elderly health improves or because retirement becomes more enjoyable (as Costa (1998) suggests has happened historically) then the risk premium increases. A low risk job guarantees individuals their "old" utility. If expected utility when retired is high, individuals will need more compensation to take a gamble when young.

We estimate value of life from measured labor market compensating differentials for risk taking using the 1940-1980 censuses and fatality data from the Bureau of Labor Statistics (see the Data Appendix for details). For every census year, i we estimate single cross-sectional hedonic wage regressions of the form,

$$w_{ij} = \beta_i X_{ij} + \gamma_i f_{ij} + u_{ij} \tag{2}$$

$$\ln(w_{ij}) = \beta_i X_{ij} + \gamma_i f_{ij} + u_{ij} \tag{3}$$

where j indexes the individual, w is the hourly wage, f is the industry fatality risk (deaths per million hours worked) by 3 digit industry classification, u is the error term, and the vector X consists of age dummies indicating race, foreign birth, marital status, education, blue collar status, and residence in a metropolitan area, and state fixed effects. We restrict the sample to full-time male workers age 18 to 45, a group empirically likely to be very sensitive to risk (Viscusi and Aldy 2003), and we exclude workers in agriculture, telecommunications, transportation, and utilities and do not analyze data for 1990 or later because of non-comparability across years.[2] However, we also present results for all ages and all non-agricultural industries as a robustness check. We restrict to men because we do not have fatality rates for women by industry. The wage is trimmed in each decade by dropping the bottom and top one percent of wages. Our state fixed effects will proxy for such variables as state workers' compensation rates. Unlike most previous studies (Hersch (1998) and Viscusi and Hersch (2001) are rare exceptions), we cluster our standard errors by the 3 digit 1950 industry classification, because we are adjusting the standard errors for unobserved industry attributes (Moulton 1990). We do not control for injury rates because the data are not comparable over the years that we examine.

Figure 1 shows that in each decade and for every broad industry category job safety has been rising. Differences in job fatality risk between mining, construction, manufacturing, and other industries (mainly trade) narrowed since 1940, with most of the decline occurring before 1970. In 1940 fatality rates in the two most dangerous industries, mining and construction were 6 and 10 times, respectively, greater than those in manufacturing. By 1980 fatality rates in mining and construction were only 2 and 3 times, respectively, as high as those in manufacturing.

Estimating the hedonic wage regression at five points in time yields five estimates of the marginal valuation of risk. A hedonic identification issue arises as to how to interpret changes in these compensating differentials over time. Given that job safety is rising over time, if we observe that compensating differentials are rising then this is evidence that the demand for job safety must be rising. If workers experienced diminishing returns from job safety then safety improvements could have led to declining compensating differentials over time.

Our cross-sectional hedonic wage regressions will recover the value of life provided that labor markets are perfectly competitive, that observables control for differences in productivity,

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[2] As discussed in the Data Appendix, the BLS changed its survey methods in 1992 and no longer provide comparable data by industry for 1990.

and that all workers have the same preferences over money and risk and are informed about actuarial risk probabilities. A recent literature has examined how relaxing these conditions affects the interpretation of hedonic wage regression estimates. Within each year several factors will lead us to underestimate the true value of life. Because we observe those with the lowest value of life in risky jobs, we will underestimate willingness to pay for safety. If workers are heterogeneous in their abilities and if the econometrician can only partially observe workers' skills, then we will underestimate the value of life in any given year. According to Hwang, Reed, and Hubbard's (1992) derivation, in the limit the ratio of the estimated fatal injury coefficient, $\hat{\gamma}$, to the true fatal injury coefficient, γ, is

$$\text{plim}(\frac{\hat{\gamma}}{\gamma}) = \frac{\tau\omega^2 - \delta(1-\tau)\omega(1-\omega)}{\tau\omega^2 + \delta(1-\tau)(1-\omega)^2}, \tag{4}$$

where ω is the average share of renumeration taken in the form of wages instead of benefits or ameneties, τ is the proportion of wage dispersion due to workers' differing tastes, and δ is the degree of unobserved productivity heterogeneity which is equal to zero when workers' productivity is perfectly observed. In this case the bias is equal to zero. When δ is big enough, i.e. $\delta > \tau\omega/[(1-\tau)(1-\omega)]$, then the estimated coefficient on fatality rates will have an incorrect negative sign. If all renumeration were taken in the form of wages then the bias would be zero. It would also be zero if all of wage dispersion were due to workers' differing tastes, rather than to their productivity.

Other factors will bias our estimates of the value of life upwards. Shogren and Stamland (2002) argue that once one accounts for both risk preference and the personal ability to reduce risk or death or injury, then value of life estimates are biased upwards even if workers self-select their jobs on the basis of their value of life and their skill. However, if workers' safety related productivity is observed by employers workers will settle into separate labor market equilibria and there will be no bias (Viscusi and Hersch 2001; Viscusi forthcoming). A second reason why our estimates of value of life may be biased upwards in each year is because the BLS data over-sample large firms and therefore underestimate risk. A third reason for upward bias is because aggregation is at the industry level rather than the worksite level (LaLive forthcoming). Also, if older workers have a lower value of life our age restriction will yield an upward bias in each year. In addition, because we cannot control for injury rates, this biases our estimates of the value of life upwards.

We view our trend as yielding a lower bound estimate of changes in the value of life. According to the Hwang, Reed, and Hubbard's (1992) bias formula (Equation 4), the downward bias will be greater in 1980 than in 1940 even if there were no change in workers' unobserved productivity because benefits have become a bigger share of renumeration. We suspect that the proportion of wage dispersion due to tastes has always been low. The degree of unobserved worker productivity was probably high in both 1940 and 1980, but it is most likely to be higher in recent data, so again the bias in 1940 will be lower than in 1980. Under Shogren and Stamland's (2002) conditions, our estimate of the trend in the value of life will be a lower bound if in the past the marginal worker was more likely to be low skilled in reducing his personal fatality rate, perhaps because knowledge was less likely to diffuse to marginal workers.[3] If the omitted variables bias from excluding injury rates was greater in 1940 then in 1980 because injury rates were higher in 1940, then we will underestimate the increase in the value of life. Because jobs have become safer, then if worker utility is concave with respect to safety, our linear specification will yield a declining estimate of the value of life. If the proportion of risk lovers is constant over time in the working population, then given that the share of jobs in risky occupations has shrunk,

[3]Our estimate of the trend in value of life will be an upper bound if today the marginal worker is more likely to be relatively low skilled in coping with job risk, a plausible scenario only if the high skill workers have been able to process increases in knowledge.

the marginal worker would require a lower risk premium. Because workers' compensation now replaces a higher percentage of after-tax income than it did in 1940, compensating differentials are now lower. Therefore these factors will impart a downward bias to our estimate of trends in the value of life.

3 Trends

Average risk can fall either because of composition shifts of workers into less dangerous industries or because of within industry risk declines. Table 1 shows that declines in industry specific fatality risk, not the shift of workers into less dangerous industries are the main cause for the decline in overall job risk. Eighty-two to 100 percent of the decline in deaths per million hours worked is due to within industry declines, with declines in construction and in manufacturing accounting for the majority of the change. Given that rising globalization has permitted the export of the more dangerous jobs abroad, it is striking that declining fatality rates within industry and not the shift of workers out of industries has lowered overall fatality rates.

Several factors probably accounted for the decline in fatality rates. These include exogeneous upgrading of the capital stock, firms' profit maximizing response to compensating differentials, union emphasis on job safety and high unionization rates in the 1940s and 1950s, increases in state workmens' compensation payments in the 1940s and 1950s, workers' greater safety precautions as firms moved away from piece rates, and government regulation. Until 1960 regulation was primarily on the state level, although the Fair Labor Standards Act of 1938 created the Division of Labor Standards to help state governments improve their job safety laws and their administration. In 1960 the Labor Department imposed federal standards on the states using its powers under the Walsh-Healey Act of 1936 to bar employment under hazardous conditions. Federal oversight increased with the 1970 Occupational Safety and Health Act (OSHA). Using longitudinal firm data, Gray and Jones (1991) report that OSHA health inspections increased compliance with safety legislation. Viscusi (1992), however, finds only an impact in the case of cotton dust. Relatively little work has been done on the effect of earlier legislation. Lewis-Beck and Alford (1980) find that federal coal mining safety regulations enacted in 1941 and 1969 reduced fatality rates. Fishback (1986) finds no effect of state mining laws enacted in the early 1900s on fatal accident rates in coal mining.

How have the socioeconomic and demographic characteristics of who works in risky industries changed? Average risk exposure between men and women narrowed (see Figure 2). In 1940 men's average risk exposure was almost 4 times greater than that of women and in 1980 it was twice as high. Because we do not have industry fatality data by sex, this trend represents a composition shift.

Table 2 illustrates changes in risk incidence with results from regressions for men in each decade, i, of the form

$$f_{ij} = \beta_i X_{ij} + u_{ij} \; , \tag{5}$$

where f is the fatality rate in individual j's industry (deaths per million hours worked), X is a vector of socioeconomic and demographic characteristics, of broad industry dummies, and of state fixed effects, and u is an error term. Controlling for broad industry group (mining, construction, manufacturing, trade, and other), risk exposure by race, education, and age has become much more equal over time. The difference in fatality rates faced by non-whites and whites controlling for other characteristics was 0.041 in 1940 and zero in 1980.[4] The difference in fatality rates faced by those with less than a high school education compared to those with a high school education

[4]However in recent data fatality rates by race show that blacks' fatality rates are somewhat higher (Viscusi forthcoming).

was 0.014 in 1940 and 0.003 in 1980. Differences in fatality rates between those with less than a high school education and the college educated are similar. Each year of age added roughly 0.004 to the fatality rate in 1940 but nothing thereafter. Interestingly, the foreign-born were not exposed to more risk.

4 Results

4.1 Hedonic Regressions

How has the wage-risk relationship changed over time? Figure 3 presents non-parametric regressions illustrating that the wage (adjusted for mean state wage) increased with industry risk in both 1940 and 1980 among men with some high school education.[5] The continuous increase in wages as fatality rates rise implies that our results do not depend upon one or two high risk and high wage industries and that our identification does not come from extreme risk. This figure also shows that in 1980 the relationship between wage and risk was steeper than in 1940, suggesting that risk compensation has increased over time. To control for other characteristics that may affect the wage-risk relationship we turn to the ordinary least squares regressions specified in Equations 2 and 3.

Our ordinary least squares regressions, under various sample specifications and using various sets of control variables, all point to rising risk compensation over time. Tables 3 and 4 present results from our preferred linear and log-linear hedonic regressions, respectively, controlling for race, foreign-birth, marital status, education, blue collar status, residence in a metropolitan area, and state fixed effects. In both cases we stratify by ages 18-30 and 31-45.

[5] We adjusted the wage by subtracting mean wage within the state from workers' wages. Our non-parametric regressions use a Gaussian kernel and Nadaraya-Watson kernel smoother with a bandwith of 0.4 (Härdle 1991: 25, 147-89).

We believe that the risk premium will be more pronounced for younger workers to entice them to enter dangerous jobs, particularly when younger workers may not have enough experience to reduce their own personal fatality rate, even though their reflexes may be better.[6] We present results for all educational levels and also for some high school or high school graduates, groups that are likely to be exposed to job risk within an industry.

Our linear wage specification shows a clear secular increase in the job risk premium. For those with some high school or high school graduates, the coefficient on the fatality rate is significantly different from zero in all years except for 1960 in both the linear and the log-linear specifications. Although we lose significance in other stratifications of the data, particularly in 1970, the magnitude of the coefficients suggests an upward trend. (Because we are clustering our standard errors by industry, we are reducing the statistical significance of our fatality rate coefficients. Moulton (1990) shows that t-statistics fall dramatically with clustering. When we calculate White standard errors to correct for heteroskedasticity and do not cluster all of our coefficients are statistically significant at the 0.001 percent level.) We also find that compensation for job risk is generally greater for our younger workers.[7] A behavioral explanation would be that workers need to be offered a large initial salary to enter the field, but then have flat experience profiles on the job as they become inured to danger.

Table 5 presents coefficient estimates for all men age 18-64 in all non-agricultural industries (including our previously excluded categories of transportation, utilities, and communications). Although the standard error increases, our point estimates are similar and indicate an upward trend. (When we do not cluster our standard errors by industry all coefficients are statistically significant.) We also tested our specification by including a quadratic term for fatality

[6]Workers older than 45 will have the experience but will not have the reflexes or the strength.

[7]As noted above, this may reflect measurement error.

risk in the estimating equation. This term was insignificant for all years except 1980 but the point estimates suggested that the return to risk is decreasing in risk.

We controlled for non-fatal accident risk to investigate whether our estimates of the value of life within a single year will be biased upward even though our measures of non-fatal accident were not comparable across years. Prior to 1970 nonfatal accident rates were so highly collinear with fatality risk that the estimated coefficient on nonfatal accident rates was negative. In 1980 the coefficient on fatality risk was positive and statistically significant. Controlling for fatality risk in our first specification in Table 3 (see the first line in the Table) reduces the coefficient on the fatality rate in 1980 from 5.662 to 4.205 ($\hat{\sigma}$=1.862, with clustering), suggesting that omitting nonfatal accident risk biases upward our estimate of value of life in each year. Note that a decrease of roughly 25 percent in the coefficient on the occupational mortality rate after accounting for non-fatal injury risk is almost identical to the percentage decline found by Cousineau et al. (1992) for Canada.

We ran several robustness checks in which we examined other sample restrictions and additional controls. When we restricted to the college-educated, a group unlikely to be in dangerous jobs even controlling for industry, none of our coefficients on injury rates were statistically significantly different from zero, as expected. We pooled the data and included state and industry fixed effects. This yielded coefficients of similar magnitude, but only the coefficients on the year dummies for 1950 and 1980 were statistically significantly different from the coefficient on the year 1940 dummy. We ran our regressions without miners, a group that Figure 1 showed experienced the most dramatic declines in injury rates. For high school or high school graduates age 18-30, this left the 1940 coefficient virtually unchanged, the 1950 coefficient smaller and no longer statistically significant, the 1960 and 1970 coefficients larger and the latter statistically significant, and the 1980 coefficient smaller but still statistically significant. We ran our specifications controlling for such industry characteristics as average age, percent high school, some college, and college-educated, percent female, percent non-white, and percent foreign-born. The magnitude of our coefficients was smaller and significance fell, but the trend was still there. We investigated using industry pollution (as proxied by energy intensity) as a control for job amenities. Our coefficient on energy intensity was not statistically significant and our other coefficients were unaffected. We also directly controlled for expected workers compensation benefits based upon individual earnings and state formulas and for state union rates, but this yielded very similar results to those presented in our tables because our state fixed effects controlled for these differences.

We investigated using wife's earnings as an instrument for the fatality risk faced by married men because men whose families have greater earnings capacity are likely to choose safer jobs (Garen 1988). Our instrument was only weakly correlated with industry fatality rates and the coefficients on our instrumented fatality rate were many orders of magnitude larger than those on our uninstrumented fatality rate.

We recognize that discrete choice methods such as those presented in Berry, Levinsohn, and Pakes (1995) could be used be used to more directly estimate whether marginal valuations for safety have increased over time. A discrete choice demand study would face at least two challenges. Because there are 48 states and over 100 industries, the dimensionality of this problem would be cumbersome. In addition, as documented in industrial organization studies such as Berry, Levinsohn, and Pakes (1995) and Petrin (2002), a discrete choice model of risky industry choice would face the extra challenge of instrumenting for wages. The industry wage would be an explanatory variable in a discrete choice model. It is likely to be correlated with observed industry attributes embedded in the error term. In a hedonic model, industry wage is a dependent variable and ordinary least squares will yield consistent implicit price estimates if the industry risk level is uncorrelated with the error term.

4.2 Value of Life

Table 6 shows that value of life has increased by 300 to 400 percent between 1940 and 1980, rising from roughly 1 million 1990 dollars in 1940 to 4 to 5 million 1990 dollars in 1980.[8] Using our value of life estimates for workers with some high school and regressing the logarithm of value of life on the logarithm of per capita GNP we found that the elasticity of value of life with respect to per capita GNP was 1.5 (95% CI=[0.9, 2.1]) using the log-linear specification and 1.7 (95% CI=[1.3, 2.1]) using the linear specification.[9] When we controlled for average fatality risk faced by the men in our sample our estimated elasticity rose to 2.07 (95% CI=[1.5, 2.6]) in the linear specification and to 2.03 (95% CI=[0.6, 3.5]) in the log-linear specification. When we regressed the logarithm of the value of life on the logarithm of the hourly wage of men in our sample rather than on the logarithm of per capita GNP our estimated elasticities were similar, but the confidence intervals were very much wider. Using value of life derived from our linear specification our estimated elasticity was 1.7 (95% CI=[0.3, 3.1]) and using value of life derived from our log-linear specification our estimated elasticity was 1.4 (95% CI=[-0.06, 3.0]). When we regressed on the logarithm of per capita GNP our estimated value of life derived from the sample of men age 18-64 in all non-agricultural industries using the linear specification our estimated elasticity increased slightly to 1.8 (95% CI=[0.7, 2.8]).

How do our estimates of value of life and of the income elasticity of value of life compare with those of other researchers? Table 7 lists value of life estimates from other studies in 1990 dollars. Our estimates for 1970 of 3 to 5 million 1990 dollars are on the low end of Viscusi's (1978) estimates of 3 to 8 million dollars for 1969-1970. Our estimates for 1980 of 4 to

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[8] For our linear specification, value of life is $\gamma \times 10^6$ and for our log-linear specification it is $\gamma \times w \times 10^6$ where γ is the coefficient on fatality risk and w is the hourly wage.

[9] Even if we exclude the 1980 data on the grounds that the change in the reporting system in the 1970s biases upward our 1980 value of life estimate, we still find that our estimated elasticity is 1.3 (95% CI=[0.3, 2.3]) using the log-linear specification and 1.6 (95% CI=[0.9, 2.2]) using the linear specification.

5 million dollars are within the range of 3 to 6 million found by Viscusi (1993) for data from the 1980s. Our 1940 estimate of 1 million 1990 dollars is similar to that observed in Taiwan in 1987 and to Hong Kong in 1991 (Hammitt et al. 2000; Siebert and Wei 1998).

Our elasticity of value of life with respect to per capita GNP of 1.5 to 1.7 is higher than that derived from most meta-analyses but lower than that obtained from individual level fatality risk and wage data. For example, Viscusi and Aldy's (2003) meta-analysis yields an income elasticity of 0.5-0.6; Mrozek and Taylor's meta-analysis yields an income elasticity of 0.4-0.5; Miller's meta-analysis yields an income elasticity of 0.9-1.0; and Liu, Hammit, and Liu's meta-analysis yields an estimate of 0.5. Bowland and Beghin's (2001)'s meta-analysis which finds an income elasticity of 1.7-2.3 is the exception. Using experimental data on individuals Viscusi and Evans (1990) obtain an income elasticity of 0.6-1.0. Hammitt et al. (2000) used individual level data for Taiwan from 1982 to 1997 to obtain an income elasticity of value of life with respect to per capita GNP of 1.5 to 2.5. The large variation in estimated income elasticities in different meta-analyses may reflect differences in sample construction (Viscusi and Aldy 2003). We do not know why income elasticities estimated from individual level data are higher than those estimated from meta-analyses but suspect that here too differences in sample construction may play a role.

Using our elasticity of per capita GNP with respect to per capita GNP of either 1.5 or 1.7 we estimate that the value of life in 1900 was 0.3 million dollars and in 1920 was 0.6 million dollars. Our estimated value of life in 1900 is therefore greater than that found by Fishback and Kim (1993) within railroads in the United States circa 1900 and our estimated value of life in 1920 is within Fishback's (1992) estimate of 0.2 to 0.8 million for coal mining in the United States in the 1910s and 1920s.[10] We predict that in 2000 the value of life was 6.1 to 8.7 million 1990 dollars or 8.5 to 12.1 2002 dollars. Note that our estimate is higher than Viscusi and Aldy's

[10]Because coal and railroads were both high risk industries, value of life estimates should be lower than for the population as a whole.

(2003) preferred estimate of 7 million dollars, but within the range of 5 to 12 million found by half of the studies done using U.S. data (see Viscusi and Aldy 2003).

5 Implications for Mortality Declines

One of the implications of rising value of life is that evaluations of the benefits of mortality reductions or health improvements over the entire twentieth century underestimate the value of current improvements in health or mortality relative to the dramatic improvements of the past. Table 8 illustrates. The largest age-adjusted mortality declines occurred prior to 1960 when infant mortality fell sharply. In 1900 the probability of a child dying before age 1 was 0.15 and his life expectancy at birth was 48. By 1960 the probability of death before age 1 was 0.03 and life expectancy at birth was 70. Applying our estimated values of life for 1980 to all years, we would conclude that the biggest gains were indeed prior to 1960 when the biggest change in quantities occurred. If we assumed that the income elasticity of value of life was equal to one and used the ratio of 1980 value of life to 1980 GNP to calculate values of life, then we would conclude that the value of mortality declines was lower between 1980 and 2000 than between 1940 and 1960. Using our calculated and predicted value of life estimates from our specifications demonstrates that the largest gains occurred after 1960 when mortality gains were relatively marginal and that gains between 1980 and 2000 were substantial.[11] After 1960, the largest increases in life expectancy came in older age mortality. Life expectancy at age 65 increased by 3 years over the 60 years between 1900 and 1960, but by another 3 years over the shorter time span between 1960 and 2000. Under the constant, income elasicity of one, and income elasticity greater than one value

[11] For simplicity, we value life at all points in the life-cycle equally. Although many studies (e.g. Murphy and Topel 2003) follow Rosen (1988) in assuming that the value of of life first rises and then decreases with age, Johansson (2002) shows that this is not necessarily true and depends upon the path of consumption over the life-cycle.

of life scenarios, per capita national health care expenditures are increasing faster than the value of mortality declines, but when we do not allow for actual changes in the value of life, our health care sector seems much less productive.

6 Conclusion

Studies from the United States past and from developing countries suggest that value of life increases with economic development, but because these studies are not comparable in either their sample selection or their specifications, calculating an income elasticity from these value of life estimates is problematic. Estimating comparable regressions from 1940 to 1980, this paper found that value of life in the United States has indeed been rising and that the estimated elasticity of value of life with respect to per capita GNP was 1.5 to 1.7. The growing premium that workers have been receiving for working in unsafe jobs and the simultaneous increase in job safety implies that as labor demand shifts workers are not paying for increases in job safety, though of course wages would be even higher if there were no increase in safety.

Our findings on rising value of life have implications for evaluating current health and safety regulations. For investments where the benefits accrue for several decades, using current value of life estimates will underestimate the economic gains. This will be particularly true for developing countries that are growing rapidly. Although the Environmental Protection Agency in the United States adjusts value of life for income growth, their use of an income elasticity of 0.4 suggests that they are still underestimating the economic gains to health and safety.

Our findings also bear on prospective evaluations of the benefits of mortality reductions or health improvements over the entire twentieth century. Using a constant value of life will underestimate the value of current improvements in health or mortality relative to the dramatic improvements of the past. Assuming an income elasticity of value of life that is equal to one will underestimate the value of the most recent improvements in mortality and paint our health care sector as less productive than it really is. We showed that accurately allowing for changes in prices, the largest benefits of improved mortality occur after 1960 and are substantial even during the last twenty years when mortality gains were relatively marginal compared to those of the first half of the twentieth century.

Data Appendix

We use the 1940-1980 censuses of population and housing.[12] We link by 3 digit industry codes to fatality data published by the Bureau of Labor Statistics (BLS). We restrict our sample to full-time male workers age 18 to 45.

1. **Industry fatality rates**. Fatality rates are per million hours worked. Rates for 1940 are from *Industrial-Injury Statistics by States, BLS Bulletin No. 700* and for mining from *Minerals Yearbook, 1941*, those for 1950 from *Work Injuries in the United States During 1950, BLS Bulletin No. 1098*, those for 1960 are from *Injury Rates By Industry, 1963, BLS Report No. 295* (earlier years do not report fatality rates), for 1970 from *Injury Rates By Industry, 1970, BLS Report No. 406*, and for 1980 from Appendix E in Leigh (1995). The data provided by Leigh (1995) are from unpublished BLS tabulations. They are per 100,000 workers. All rates were converted to million hours worked using data from *Hours, Wages, and Earnings in the United States*. We were not able to obtain fatality data by industry for 1990 because the BLS no longer provide these data and, although they could be re-estimated from their records, the cost was prohibitive. In 1992, the BLS changed its survey methods, including smaller establishments, workers on small farms, the self-employed and family workers, and

[12] We use the integrated public use micro samples available at http://www.ipums.umn.edu/.

public sector workers. Reported fatality rates rose.

The BLS obtained fatality rates by surveying firms. Because large firms are over-sampled in 1940-1970, accident rates are underestimated. Firms faced fines for non-reporting only beginning in 1972. The BLS also reported injury rates. These are for injuries resulting in permanent partial disability and temporary total disability. After 1971 the definition of an injury changed to those established under the Williams-Steiger Occupational Safety and Health Act of 1970. This change expanded the definition of an injury. For details see the reports cited above and also various issues of *Handbook of Labor Statistics*.

We do not have fatality information for agricultural workers and we excluded fatality information for workers in telecommunications, transportation, and utilities. When we examined the fatality data for workers in telecommunications, transportation, and utilities, we found considerable swings across decades. We suspect that this was due both to changes in industry definitions and to relatively few of these types of firms reporting. After these exclusions we were left with 163 industries in the census microsamples and we were able to link 110 of these industries to fatality data. We do not limit ourselves to a balanced panel of industries. When we do the magnitude of our coefficients on fatality risk remains unchanged but the standard error increases. The percentage of 18-45 year-old men in manufacturing in our decade samples ranges from 47 to 55 percent. Our next largest category is "other than manufacturing, mining, or construction," ranging from 39 to 26 percent of our sample.

2. **Wage**. Our hourly wage variable is constructed from annual wage and salary divided by annual hours worked (current weekly hours multiplied by weeks worked in the past year). For 1960 and 1970, where we only have intervalled hours and weeks data, we take the midpoint. We adjust all wages to 1990 dollars. We multiply the topcode by 1.45. In 1980 a larger proportion of the population was covered by topcoding than in 1970. However, our

results remain unchanged when we impose a new topcode in 1970. We trim the bottom and top one percent of wages.

3. **Race**. Our race variable is a dummy equal to one if the worker was not white.

4. **Foreign birth**. Our foreign birth variable is a dummy equal to one if the worker was born abroad.

5. **Marital status**. Our marital status variable is a dummy equal to one if the worker was married.

6. **Education**. We create four dummy variables for education – less than high school, some high school or high school graduate, some college, and college graduate.

7. Blue collar. Our blue collar variable is a dummy equal to one if the

8. Metropolitan area resident. This is a dummy variable equal to one if the worker lived in a metropolitan area, as defined by that year's census. Definitions vary across years.

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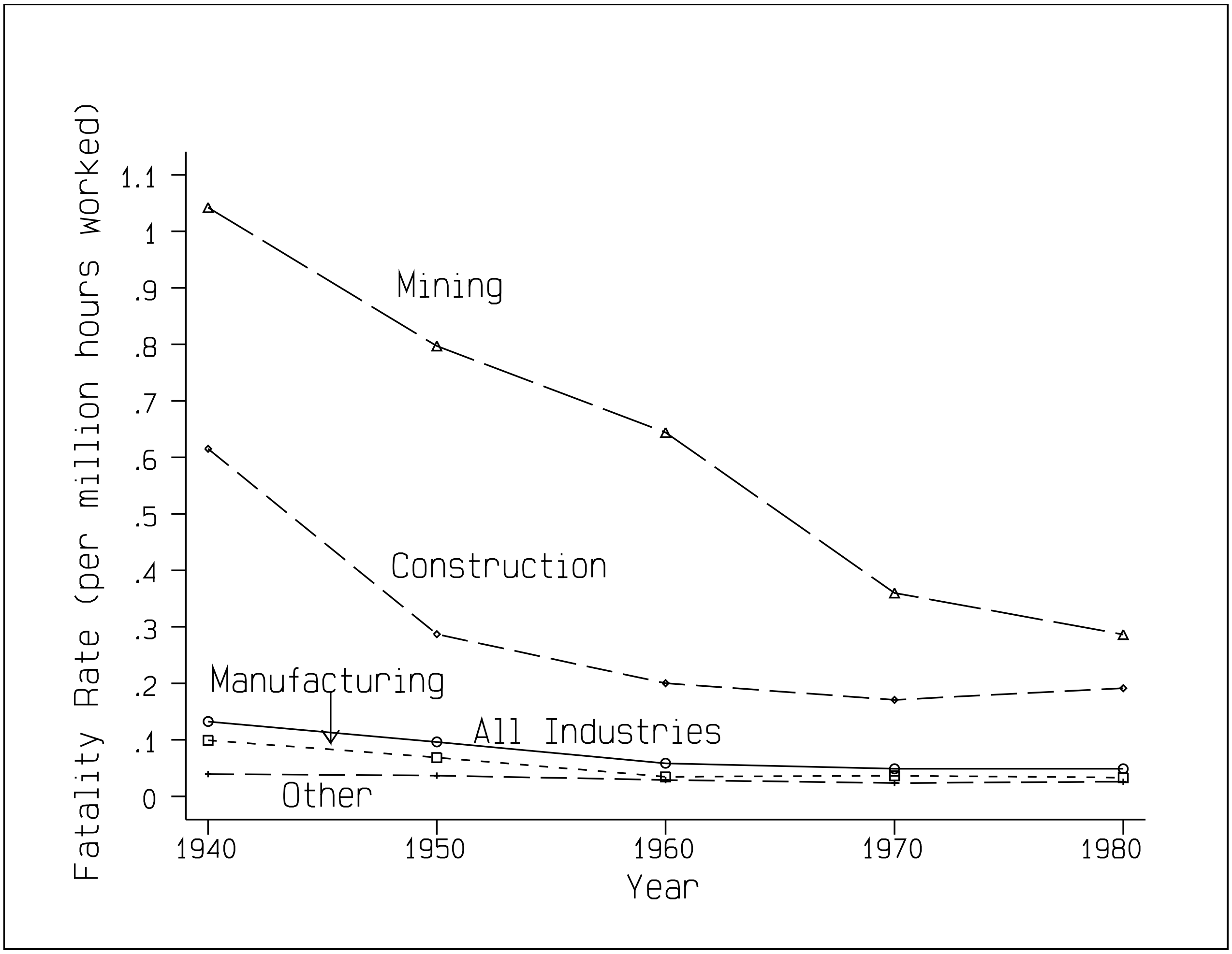
Table 1: Decomposition of Decline in Deaths per Million Hours into Within Industry Decline and Declines Due to Shifts in Workers Across Industrial Sectors, 1940-1980

	Sample		Weighted Sample	
	Δ	% Δ due to	Δ	% Δ due to
Decline in deaths per million hours worked 1940-1980	0.0835		0.0892	
Decline due to shifts in workers across industries	-0.0033	-3.8%	0.0158	17.7%
Decline due to within industry risk declines	0.0867	103.8	0.0733	82.3
Decline due to manufacturing	0.0270	32.3	0.0229	25.7
Decline due to mining	0.0159	19.1	0.0135	15.1
Decline due to construction	0.0370	44.3	0.0312	35.0
Decline due to trade	0.0065	7.8	0.0055	6.1
Decline due to other industries	0.0002	0.3	0.0003	0.4

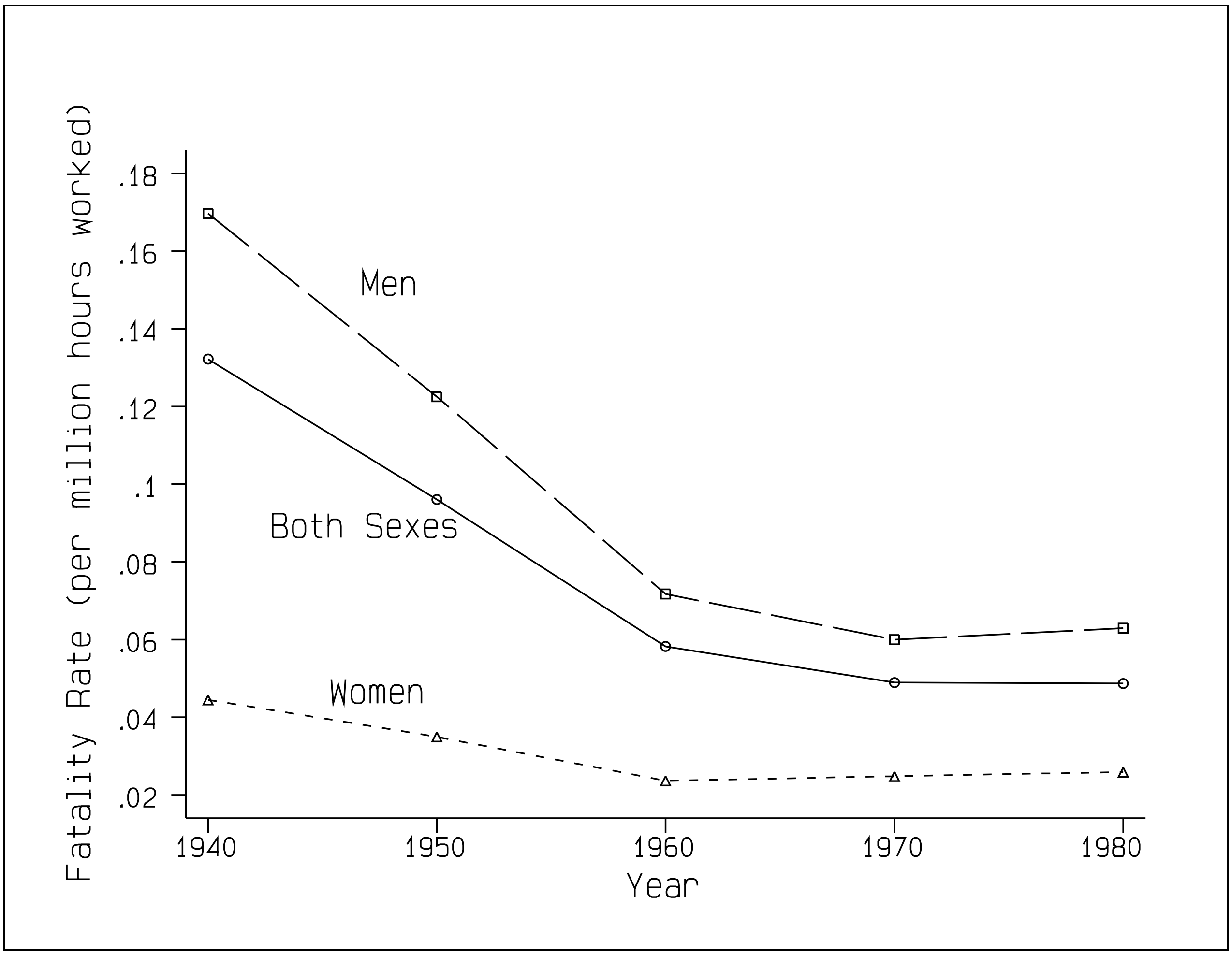
Note: Estimated from the Integrated Public Use Census Samples for both sexes and BLS death rates. The industries are manufacturing, mining, construction, trade, and other. Other does not include agriculture, communications, transportation, and utilities. The weighted sample assumes that 3 digit industries for which risk levels are unknown have the same risk of death as the broad industry average. The sample is restricted to full-time workers age 18 to 45.

Figure 1: Mean Industry Fatality Risk in Jobs Employing Workers Age 18 to 45, 1940-1980



Source: Estimated from the Integrated Public Use Census Samples and BLS death rates. Other industries do not include agriculture, communications, transportation, and utilities. The sample is restricted to full-time workers of both sexes. Fatality risk is measured as deaths per million hours worked for employees of all ages.

Figure 2: Mean Industry Fatality Risk in Jobs Employing Male and Female Workers Age 18 to 45, 1940-1980



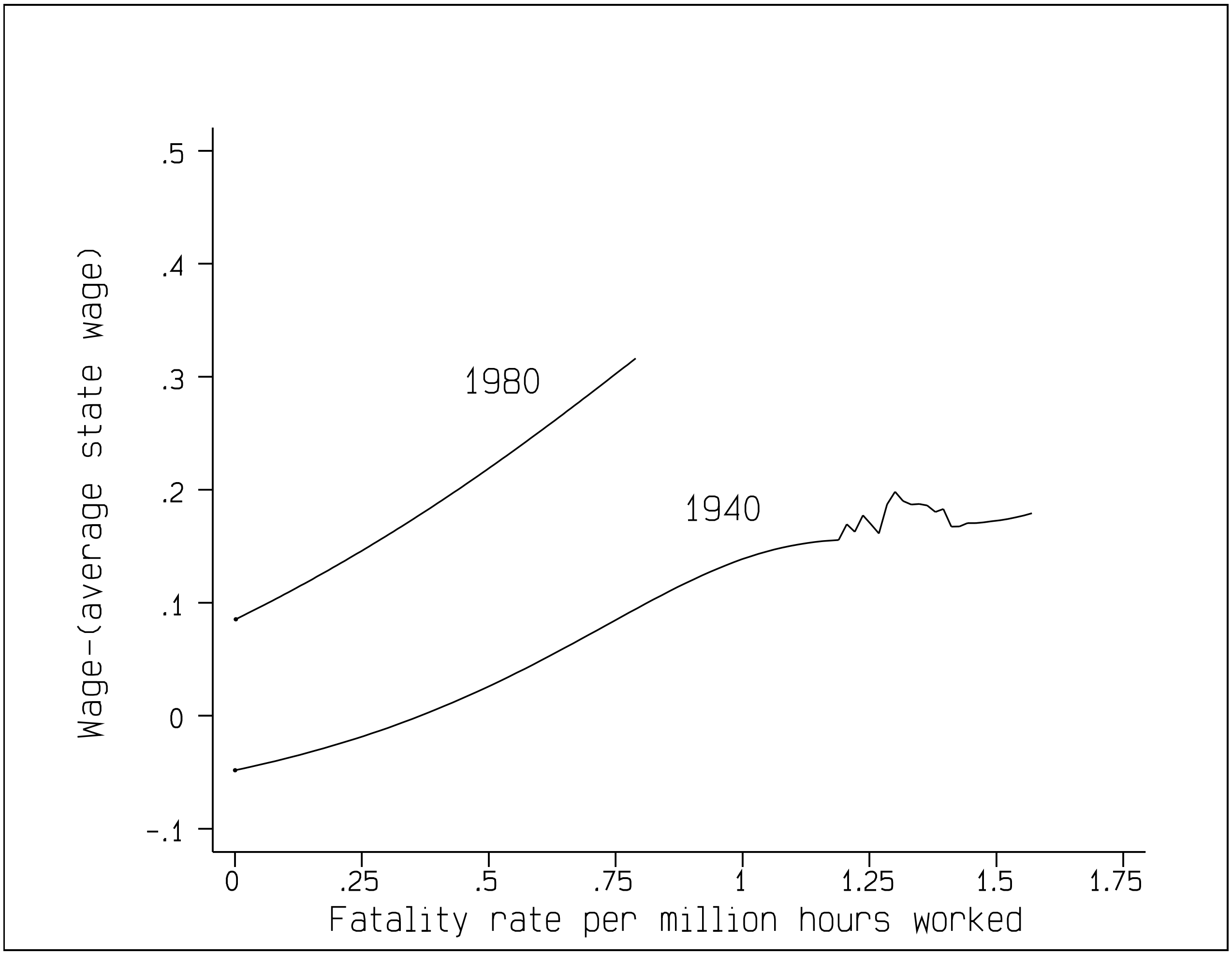
Source: Estimated from the Integrated Public Use Census Samples and BLS death rates. Workers in agriculture, communications, transportation, and utilities are excluded. The sample is restricted to full-time workers. Fatality rates are deaths per million hours worked for employees of all ages and for both sexes.Fatality rates are deaths per million hours worked for employees of all ages and for both sexes.

Table 2: Characteristics of Men in High Risk Industries Controlling for Socio-economic and Demographic Characteristics and for Industrial Sector, 1940-1980

	1940	1950	1960	1970	1980
Dummy=1 if					
non-white	0.041‡	0.029‡	0.014‡	0.008‡	0.000
	(0.002)	(0.002)	(0.001)	(0.001)	(0.000)
foreign-born	-0.002	-0.006†	-0.005‡	-0.003‡	-0.004‡
	(0.002)	(0.003)	(0.001)	(0.001)	(0.000)
married	-0.043‡	-0.010‡	-0.002†	0.000	0.001‡
	(0.001)	(0.002)	(0.001)	(0.001)	(0.000)
Number of children	0.007‡	0.002‡	0.001†	0.001‡	0.001‡
	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
Dummy=1 if education					
less than high school					
high school	-0.014‡	-0.008‡	-0.006‡	-0.007‡	-0.003‡
	(0.001)	(0.002)	(0.001)	(0.001)	(0.000)
some college	-0.016‡	-0.009‡	-0.005‡	-0.007‡	-0.005‡
	(0.002)	(0.003)	(0.001)	(0.001)	(0.000)
college	-0.011‡	-0.019‡	-0.009‡	-0.008‡	-0.004‡
	(0.002)	(0.003)	(0.001)	(0.001)	(0.000)
Age	0.004‡	0.000	0.000	0.000	0.001‡
	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)
Age squared	0.000‡	0.000	0.000	0.000	0.000‡
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Dummy=1 if					
blue-collar	-0.007‡	0.018‡	0.004‡	0.006‡	0.006‡
	(0.001)	(0.002)	(0.001)	(0.001)	(0.000)
in metropolitan area	-0.018‡	-0.019‡	-0.007‡	-0.006‡	-0.002‡
	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Dummy=1 if industry					
trade					
manufacturing	0.041‡	0.042‡	-0.008‡	0.009‡	-0.010‡
	(0.001)	(0.002)	(0.001)	(0.001)	(0.000)
mining	0.724‡	0.737‡	0.559‡	0.334‡	0.238‡
	(0.003)	(0.004)	(0.002)	(0.001)	(0.001)
construction	0.536‡	0.244‡	0.151‡	0.134‡	0.144‡
	(0.002)	(0.002)	(0.001)	(0.001)	(0.000)
other	-0.059‡	0.039‡	-0.023‡	-0.007‡	-0.027‡
	(0.002)	(0.002)	(0.001)	(0.001)	(0.000)
Constant	0.051‡	0.030‡	0.044‡	0.029‡	0.034‡
	(0.010)	(0.013)	(0.004)	(0.004)	(0.002)
Adjusted R^2	0.728	0.649	0.637	0.508	0.685
Number of observations	90,050	10,895	110,225	107,406	176,507

Note: Estimated from the Integrated Public Use Census Samples and BLS death rates. The dependent variable is the industry fatality rate per million hours worked. See Equation 5 in the text. Other industry does not include agriculture, communications, transportation, and utilities. The sample is restricted to full-time workers age 18 to 45. All regressions include state fixed effects. Robust standard errors, clustered on 3 digit industry, in parentheses. The symbols † and ‡ indicate that the coefficient is significantly different from 0 at the 5 and 1 percent level, respectively.

Figure 3: Non-Parametric Value of Life Regressions



Source: Estimated from the Integrated Public Use Census Samples and BLS death rates. Results are from a kernel regression with Gaussian kernel and the Nadaraya-Watson kernel smoother with a bandwidth of 0.4 (Härdle 1991: 25, 147-89). The sample is restricted to full-time male workers age 18-45 with some high school education and excludes workers in agriculture, telecommunications, transportation, and utilities. Mean wage within the state was substracted from workers' wages.

Table 3: Coefficients on Deaths per Million Hours Worked When the Dependent Variable is the Hourly Wage, 1940-1980

	1940	1950	1960	1970	1980
All educational levels					
Age 18-30					
Coefficient	0.802	1.549	1.604	3.301	5.662
	(0.273)‡	(0.709)†	(1.277)	(2.054)	(1.763)‡
	[0.045]‡	[0.188]‡	[0.178]‡	[0.246]‡	[0.223]‡
Adjusted R^2	0.275	0.216	0.279	0.264	0.223
Observations	44,340	12,967	42,689	48,541	93,302
Age 31-45					
Coefficient	0.982	1.302	1.658	3.293	5.033
	(0.314)‡	(0.628)†	(1.160)	(2.902)	(2.214)†
	[0.058]‡	[0.175]‡	[0.142]‡	[0.231]‡	[0.291]‡
Adjusted R^2	0.261	0.199	0.265	0.290	0.216
Observations	45,710	16,213	67,536	58,865	83,205
Some high school or high school graduate					
Age 18-30					
Coefficient	0.996‡	1.755‡	2.086	3.474*	5.347‡
	(0.237)‡	(0.622)‡	(1.377)	(2.074)*	(1.736)‡
	[0.068]‡	[0.265]‡	[0.237]‡	[0.292]‡	[0.264]‡
Adjusted R^2	0.233	0.178	0.248	0.218	0.220
Observations	24,936	7,745	25,561	32,082	55,749
Age 31-45					
Coefficient	1.167‡	1.361*	2.132	4.397	5.259†
	(0.388)‡	(0.717)*	(1.416)	(3.804)	(2.642)†
	[0.117]‡	[0.292]‡	[0.214]‡	[0.336]‡	[0.371]‡
Adjusted R^2	0.174	0.100	0.132	0.135	0.130
Observations	16,014	7,392	34,705	33,861	42,362

Note: Estimated from the Integrated Public Use Census Samples and BLS death rates using Equation 2. The reported coefficients are an estimate of γ. Workers in agriculture, communications, transportation, and utilities are excluded. The sample is restricted to full-time workers age 18 to 45. Coefficients are from an ordinary least squares regression in which the hourly wage is the dependent variables and additional independent variables are age and age squared and dummy variables for non-white, foreign-born, married, blue collar, within a metropolan area, education (less than high school, high school, some college, and college), and state fixed effects. Robust standard errors adjusted for clustering in parentheses and heteroskedasticity corrected standard errors in square brackets. The symbols $*$, †, and ‡ on the standard errors indicate that the coefficient is significantly different from 0 at the 10, 5, and 1 percent level, respectively.

Table 4: Coefficients on Deaths per Million Hours Worked When the Dependent Variable is the Logarithm of the Hourly Wage, 1940-1980

	1940	1950	1960	1970	1980
All educational levels					
Age 18-30					
	0.160	0.174	0.121	0.248	0.458
	(0.077)†	(0.108)	(0.154)	(0.155)	(0.167)‡
	[0.010]‡	[0.028]‡	[0.022]‡	[0.022]‡	[0.021]‡
Adjusted R^2	0.318	0.244	0.302	0.279	0.241
Observations	44,340	12,967	42,689	48,541	93,302
Age 31-45					
	0.157	0.118	0.099	0.193	0.296
	(0.066)†	(0.089)	(0.106)	(0.176)	(0.163)*
	[0.010]‡	[0.023]‡	[0.014]‡	[0.017]‡	[0.022]‡
Adjusted R^2	0.290	0.215	0.272	0.271	0.204
Observations	45,710	16,213	67,536	58,865	83,205
Some high school or high school graduate					
Age 18-30					
	0.220	0.203	0.191	0.271	0.436
	(0.064)‡	(0.093)†	(0.153)	(0.153)‡	(0.166)‡
	[0.015]‡	[0.038]‡	[0.028]‡	[0.026]‡	[0.025]‡
Adjusted R^2	0.274	0.205	0.275	0.251	0.239
Observations	24,936	7,745	25,561	32,082	55,749
Age 31-45					
	0.194	0.135	0.145	0.254	0.302
	(0.072)‡	(0.087)	(0.116)	(0.229)	(0.181)*
	[0.019]‡	[0.038]‡	[0.020]‡	[0.025]‡	[0.029]‡
Adjusted R^2	0.197	0.115	0.148	0.145	0.133
Observations	16,014	7,392	34,705	33,861	42,362

Note: Estimated from the Integrated Public Use Census Samples and BLS death rates using Equation 3. The reported coefficients are an estimate of γ. Workers in agriculture, communications, transportation, and utilities are excluded. The sample is restricted to full-time workers age 18 to 45. Coefficients are from an ordinary least squares regression in which the logarithm of the hourly wage is the dependent variables and additional independent variables are age and age squared and dummy variables for non-white, foreign-born, married, blue collar, within a metropolan area, education (less than high school, high school, some college, and college), and state fixed effects. Robust standard errors adjusted for clustering in parentheses and keteroskedasticity corrected standard errors in square brackets. The symbols $*$, $\dagger$, and $\ddagger$ on the standard errors indicate that the coefficient is significantly different from 0 at the 10, 5, and 1 percent level, respectively.

Table 5: Coefficients on Deaths per Million Hours Worked for Men Age 18-64 in All Non-Agricultural Industries, 1940-1980

	1940	1950	1960	1970	1980
All educational levels					
Dependent variable: hourly wage					
	0.899	1.303	1.473	3.024	5.516
	(0.303)‡	(0.695)*	(1.062)	(2.375)	(1.803)‡
	[0.034]‡	[0.108]‡	[0.088]‡	[0.134]‡	[0.155]‡
Dependent variable: log(hourly wage)					
	0.164†	0.141	0.108	0.198	0.386
	(0.069)	(0.098)	(0.105)	(0.150)	(0.149)‡
	[0.006]‡	[0.015]‡	[0.009]‡	[0.011]$^{[}\ddagger]$	[0.012]‡
Some high school or High school graduate					
Dependent variable: hourly wage					
	1.038	1.238	1.782	3.829	5.620
	(0.343)‡	(0.751)$*$	(1.254)	(2.939)	(1.871)‡
	[0.062]‡	[0.181]‡	[0.136]‡	[0.188]‡	[0.187]‡
Dependent variable: log(hourly wage)					
	0.211	0.138	0.146	0.250	0.392
	(0.072)‡	(0.098)	(0.115)	(0.183)	(0.151)‡
	[0.011]‡	[0.024]‡	[0.013]‡	[0.014]‡	[0.015]‡

Note: Estimated from the Integrated Public Use Census Samples and BLS death rates using Equations 2 and 3. The reported coefficients are an estimate of γ. Workers in agriculture are excluded. The sample is restricted to full-time workers age 18 to 64. Coefficients are from an ordinary least squares regression in which either the hourly wage or the logarithm of the hourly wage is the dependent variables and additional independent variables are age and age squared and dummy variables for non-white, foreign-born, married, blue collar, within a metropolan area, education (less than high school, high school, some college, and college), and state fixed effects. Robust standard errors adjusted for clustering in parentheses and heteroskedasticity corrected standard errors in square brackets. The symbols $*$, $\dagger$, and $\ddagger$ on the standard errors indicate that the coefficient is significantly different from 0 at the 10, 5, and 1 percent level, respectively.

Table 6: Value of Life Estimates in 1000s of 1990 Dollars, 1940-1980

	Range of estimates over all samples and specifications	Some high school or high school Graduate age 18-30 Specification: Linear	Logarithmic
1940	713-996	996	977
1950	1,122-1,755	1,755	1,340
1960	1,085-2,132	2,086	1,658
1970	2,792-4,937	3,744	2,921
1980	4,144-5,347	5,347	4,253

Note: Estimated from Tables 3 and 4. All values are in 1000s of 1990 dollars.

Table 7: Value of Life Estimates from Previous Literature in 1990 Dollars

Paper	Country Examined	Year Examined	Value of life in Millions of 1990 US $
Kim and Fishback (1993)	USA (railroads)	1893-1909	0.1
Fishback (1992)	USA (coal)	1912-1920s	0.2-0.8
Thaler and Rosen (1976)	USA	1967	0.2-2.1
Arnould and Nichols (1983)	USA	1967	0.8-9.9
Herzog and Schlottman (1990)	USA	1965	6.5-8.3
Viscusi (1978)	USA	1969-1970	3.0-8.2
Olson (1981)	USA	1973	1.5-25.7
Leigh and Folsom (1984)	USA	1974-1977	7.9-10.4
Moore and Viscusi (1988)	USA	1977	0.2-0.9
Dillingham (1985)	USA	1976-1979	0.2-7.1
Leigh (1987)	USA	1979-1984	9.6-11.4
Garen (1988)	USA	1981-1982	5.8-13.2
Leigh (1995)	USA	1980-1985	2.3-12.8
Viscusi (1993)	USA	1982-1987	2.6-6.0
Cousineau et al (1992)	Canada	1979	1.9-4.3
Martinello and Meng (1992)	Canada	1986	6.9-12.0
Marin and Psacharopoulos (1982)	UK	1975	2.6-2.9
Siebert and Wei (1994)	UK	1983	4.9-11.8
Arabsheibani and Marin (1999)	UK	1980-1985	11.2-34.4
Sandy and Elliot (1996)	UK	1986	6.8-7.9
Kniesner and Leeth (1991)	Australia	1984-85	2.3
Siebert and Wei (1998)	Hong Kong	1991	1.4
Kim and Fishback (1999)	Korea	1984-1990	0.6
Hammitt, Liu, and Liu (2000)	Taiwan	1982	0.4
		1987	1.0
		1992	3.9
		1997	3.2
Simon et al. (1999)	India	c. 1990	0.2-0.4

Note: Non-US values of life were converted to US dollars at exchange rates given by the authors or at current exchange rates.

Table 8: The Value of Mortality Declines by Period in 1000s of 1990 Dollars, 1900-2000

	1900-20	1920-40	1940-60	1960-80	1980-2000
Age-adjusted mortality decline per year per million persons	3,709	3,621	4,458	3,001	1,671
Value of annual mortality declines, per person:					
Using 1980 value of life of 5.3 million	19,832	19,361	23,837	16,046	8,935
of 4.3 million	15,774	15,400	18,960	12,763	7,106
Using income elasticity of one and 1980 value of life:					
of 5.3 million	4,792	6,420	10,753	13,195	10,721
of 4.3 million	3,765	5,044	9,491	10,446	8,476
Using average value of in each period:					
Column 3, Table 6 and predicted	1,771	2,970	6,860	11,136	11,732
Column 4, Table 6 and predicted	1,719	2,873	5,865	8,854	8,702
Increase in annual per capita national health care expenditures		102	369	1,045	1,824

Sources: Age-adjusted death rates are standardized using year 2000 standard population and are from series Hist 293, CDC/NCHS. Value of life estimates based upon an income elasticity of one calculate value of life using the ratio of 1980 value of life to 1980 GNP. Value of life estimates from columns 3 and 4 in Table 6 are from the linear (Equation 4) and log-linear (Equation 5) specifications, respectively. For 1900, 1920, and 2000 we used these specifications to predict value of life in those years, obtaining 309, 647, and 8,718 thousand dollars for these respective years using the linear specification and 317, 612, and 6,179 for these respective years using the log-linear specification. Per capita national health care expenditures are from Series B 221-235 (U.S. Bureau of the Census 1975:73) and from U.S. Centers for Medicare and Medicaid Services, Health Accounts. Per capita expenditures are only available beginning in 1929. The annual increase between 1920 and 1929 was interpolated upon 1929-1940 trends.