Monetary values of increasing life expectancy: sensitivity to shifts of the survival curve

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Abstract

Individuals' monetary values of decreases in mortality risk depend on the magnitude and timing of the risk reduction. We elicited stated preferences among three time paths of risk reduction yielding the same increase in life expectancy (decreasing risk for the next decade, subtracting a constant from or multiplying risk by a constant in all future years) and willingness to pay (WTP) for risk reductions differing in timing and life-expectancy gain. Respondents exhibited heterogeneous preferences over the alternative time paths, with almost 90 percent reporting transitive orderings. WTP is statistically significantly associated with life-expectancy gain (between about 7 and 28 days) and to a limited degree with respondents' stated preferences over the alternative time paths. Estimated value per statistical life year (VSLY) averages about $500,000, roughly consistent with conventional estimates obtained by dividing estimated value per statistical life by discounted life expectancy.

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1. Introduction

Estimates of the monetary values of changes in mortality risk are important for decision making. They are used to evaluate public policies affecting, inter alia, environmental quality, transportation safety, occupational health, and product safety (US Office of Management and Budget 2020) and can be a useful guide for individual decision making about jobs, health care, diet, exercise, and other choices (Smith and Keeney 2005).

The conventional approach to quantifying the monetary value of changes in mortality risk uses the value per statistical life (VSL; in the UK, value of a prevented fatality, VPF). VSL is defined for an individual as her marginal rate of substitution between wealth and mortality risk in a specified time period. The period is usually taken to be of short duration (often one year) and is usually the current period, though the rate of substitution between wealth and risk for longer periods (e.g., a decade) and for specified future periods have also been investigated (e.g., Krupnick et al. 2002, Alberini et al. 2004).

An alternative metric, the value per statistical life year (VSLY; in the UK, value of a life year, VOLY) is sometimes considered. VSLY is defined for an individual as her marginal rate of substitution between wealth and life expectancy (Hammitt 2013). VSL and VSLY are closely related, since a decrease in mortality risk in the current year increases life expectancy by approximately the product of the risk reduction and life expectancy conditional on surviving the current year.

More generally, any perturbation to an individual's future time path of mortality risk can be represented alternatively as a change in mortality risk over any specified period or as a change in life expectancy; hence the monetary value of the perturbation can be described alternatively as the VSL or VSLY for that change. Note that these values may depend on the time pattern of changes in mortality risk and (for VSL) on the time period over which the change in mortality risk is evaluated (Hammitt 2007, Hammitt et al. 2020).

In this paper, we seek to provide improved estimates of VSLY and how it depends on the time path of the change to the individual's mortality hazard function. As described below, the same gain in life expectancy can be produced by an infinite set of changes to the baseline hazard. In our stated-preference survey, we represent an individual's hazard function as the average annual mortality risk by decade and consider three patterns of risk reduction: a transient perturbation that decreases mortality risk for only the first decade, an additive perturbation that decreases mortality risk by subtracting a constant from the risk in every decade, and a proportional perturbation that decreases mortality risk by multiplying the risk in each decade by a constant. Setting the life-expectancy gain equal across perturbations, the probability distribution of age at death is riskier for the proportional than for the additive than for the transient perturbation (i.e., the distributions are related as mean-preserving spreads). Hence an individual who is risk-averse with respect to longevity will prefer the transient to the additive to the proportional perturbation, one who is risk-seeking will have the opposite preference ordering, and one who is risk-neutral will be indifferent among the three perturbations.

In prior work, Neilsen et al. (2010) and Hammitt and Tunçel (2015) asked survey respondents to make three pairwise choices between transient, additive, and proportional perturbations of their hazard functions producing the same increase in life expectancy. In both studies, responses were largely coherent (only 10 to 15 percent of respondents made intransitive choices and the hypothesis that respondents chose randomly can be rejected with high confidence) and there is substantial population heterogeneity in preferences. In their survey of 129 Newcastle UK residents aged about 40 years, Nielsen et al. found that the fractions whose responses are consistent with risk aversion, neutrality, and seekingness are 22, 6, and 23 percent, respectively. In a representative survey of 1024 French residents aged 20 to 69 years, the corresponding fractions are 14, 23, and 16 percent (Hammitt and Tunçel 2015). This heterogeneity implies that individuals' VSLYs should be sensitive to the associated time path of risk reduction.

In this paper, we describe and report results of a stated-preference study administered online to a representative sample of about 1,000 US residents aged 20 to 69 years. Our primary objective is to estimate VSLY and to determine how it depends on the time path of risk reduction. Respondents made pairwise choices between transient, additive, and proportional risk reductions all providing the same increase in life expectancy. Similar to prior work, we find that most respondents (88 percent) report transitive preferences but their preference orderings are diverse; the fractions of respondents whose choices are consistent with global risk aversion, neutrality, and seekingness are 12, 13, and 20 percent, respectively. We elicit WTP for risk reductions that differ in type of perturbation and life expectancy gain (ranging between about 7 and 28 days). WTP is significantly associated with life-expectancy gain and with respondents' preferences over alternative perturbations of their hazard function. On average, VSLY equals about $500,000, roughly consistent with conventional estimates obtained by dividing estimated VSL by remaining life expectancy.

In the following sections we describe the theoretical model underlying this work in Section 2, prior approaches to estimating VSL and VSLY in Section 3, the survey instrument and administration in Section 4, and the statistical models and empirical results in Section 5. Conclusions are presented in Section 6.

2. Theoretical model[1]

For an individual, let *t* denote age (equivalently, time) where *t* = 0 corresponds to the age at which the value of a reduction in future mortality risk is evaluated (so positive values of *t* denote future years). Let

f(*t*) be the probability density of dying at *t*,

s(*t*) be the survival function (the probability of not dying before *t*), and

h(*t*) be the hazard function (the probability of dying at *t* conditional on survival to *t*). Then

$$s(t) = 1 - \int_0^t f(\tau)d\tau \tag{1}$$

and

$$h(t) = \frac{f(t)}{s(t)}. \tag{2}$$

Life expectancy (remaining) is the expected number of future life-years,

$$LE = \int_0^\infty tf(t)dt = \int_0^\infty s(t)dt. \tag{3}$$

Consider an individual who evaluates mortality risks by the expected utility of longevity,

$$V = Eu(t) = \int_0^\infty f(t)u(t)dt, \tag{4}$$

where *u*(*t*) is the utility of living from age 0 to death at age *t*. The utility of a living a year at age *t* is approximately $u'(t)$ where prime denotes first derivative. The individual's utility *u*(*t*) may depend on her health, consumption, and other factors that influence wellbeing.

If *u*(*t*) is linear, the individual is risk-neutral with respect to longevity and indifferent among all survival functions with equal life expectancy. Alternatively, if *u*(*t*) is (globally) concave, she is risk-averse with respect to longevity and prefers shifts to her survival function that (holding life expectancy constant) reduce uncertainty about the time of death (i.e., shifts that "rectangularize" the survival function). If *u*(*t*) is (globally) convex, she is risk-seeking with respect to longevity and dislikes shifts that rectangularize the survival function (holding life expectancy constant).

[1] Much of this section is based on Hammitt and Tunçel (2015).

An individual's risk posture with respect to longevity can differ for values of *t* in different intervals; e.g., she could be risk-seeking for short values of longevity and risk-averse for longer values.[2] Local risk aversion, neutrality, and seekingness are characterized by whether the second derivative of the function $u(t)$ at t is negative, zero, or positive, respectively.

Preferences for longevity may depend on attributes that vary with age, such as health and income. For example, an individual might be risk-averse with respect to longevity because she anticipates diminishing marginal utility per year of life associated with age-related decreases in health. Quality-adjusted and disability-adjusted life years weight each time period by a factor that reflects health, with time lived in good health more desirable than time lived in poor health (Pliskin et al. 1980). Similarly, the utility associated with being alive at age *t* may depend on consumption of goods and services, and hence indirectly on wealth or income (Jones-Lee 1974, Hammitt 2000, 2002). An individual can influence her future health and consumption by health and financial investments (e.g., diet, exercise, retirement savings). Hence her preference for a change to her survival function may depend on past health and financial investments and the terms on which she can use annuities or other tools to manage the risk of outliving her savings (Drèze 1962, Rosen 1988).

Although it is reasonable to assume individuals are risk-averse with respect to wealth, there is no such presumption concerning risk posture with respect to longevity. Empirical evidence suggests some people are globally risk-averse, risk-neutral, or risk-seeking with respect to longevity; for others, the degree and even sign of risk aversion may depend on the duration. In a pioneering study, McNeil et al. (1978) found two (of 14 total) cancer patients were risk-seeking with respect to a few years and risk-averse with respect to longer values of longevity (the others were globally risk-averse). Delprat et al. (2015) surveyed more than 1400 Americans aged 40 to 60 years, asking them to choose between living to age 80 years for sure and binary lotteries with equal probabilities of living to ages 70 or 90, or to ages 65 or 95. The fractions of respondents whose choices are consistent with risk aversion, risk neutrality, and risk seekingness are respectively 38, 27, and 35 percent for the first lottery and 59, 22, and 19 percent for the second (riskier) lottery. Several recent studies have found predominantly risk-seeking preferences when cancer patients choose between a skewed lottery with a small probability of substantial longevity and its (small) expected value. Lakdawalla et al. (2012) found 77 percent of 150 respondents preferred a trinary lottery with a long right tail to its expected value (18 or 24 months) and Shafrin et al. (2017) found about 65 percent of

[2] This pattern could be consistent with prospect theory (Kahneman and Tversky 1979) if short values are perceived as a loss and longer values as a gain relative to the individual's reference point.

165 respondents also chose the more risky lottery over its expected value (4 years).[3] Reed et al. (2021) found 57 percent of 200 respondents preferred a binary lottery with a 0.1 probability of 10 year survival and complementary chance of imminent death to its expected value of 2 years; 36 percent chose the fixed outcome and 13 percent were indifferent.

An individual's WTP for a perturbation to her hazard function depends on the time at which she learns of the shift and her ability to adapt her plans for health and financial investment and saving to the new conditions. For small perturbations, WTP (*w*) can be approximated by the expected present value of the change in hazard multiplied by the individual's age-dependent marginal rate of substitution between income and mortality risk (VSL), i.e.,

$$w = \int_0^\infty \Delta h(t) v(t) s(t) \rho(t) dt = \int_0^\infty \Delta f(t) v(t) \rho(t) dt, \tag{5}$$

where $\Delta h(t)$ is the change in hazard, $\Delta f(t)$ is the change in marginal probability of death, $v(t)$ is minus one times the individual's age-dependent VSL, and $\rho(t)$ is her consumption-discount factor (normalized so that $\rho(0) = 1$) (Johannesson et al. 1997).[4]

Expression (5) shows that the value of a change in hazard function to an individual depends on the time path of the perturbation, the time path of her VSL (which may depend on health, income, and other factors), and how she discounts future values (e.g., the value of her consumption-discount rate, whether she discounts exponentially or hyperbolically; Frederick et al. 2002). It is intuitive that a temporary decrease of fixed magnitude in the hazard provides a greater increase of life expectancy if it occurs earlier rather than later. Nevertheless, an individual may prefer a delayed reduction in hazard if the product of her VSL, survival probability, and discount factor at the later time is larger than the corresponding product at the earlier time (Hammitt and Liu 2004).

From equation (4), the individual's change in expected utility that results from a change in her probability distribution of age at death $\Delta f(t)$ is equal to

$$\Delta V = \int_0^\infty \Delta f(t) u(t) dt. \tag{6}$$

Comparing equations (6) and (5) reveals that the product of the individual's VSL and consumption-discount factor at time *t* is proportional to her utility of living from birth to age *t*. Both functions weight the change in unconditional mortality risk $\Delta f(t)$ as a function of when it occurs. The units are

[3] Lakdawalla et al. (2012) and Shafrin et al. (2017) required respondents to express a strict preference for the lottery or its expected value; indifference (and risk neutrality) were not permitted.

[4] Note that even if $\Delta h(t) < 0$ for all t, $\Delta f(t) > 0$ for some values of *t* because $\int_0^\infty \Delta f(t) dt = 0$.

different, however. The weighting factor $v(t) \cdot \rho(t)$ in equation (5) is measured in monetary units (e.g., dollars) and the weighting factor $u(t)$ in equation (6) is measured in utility units.

The proportionality between $u(t)$ and $v(t) \cdot \rho(t)$ reveals an isomorphism between risk posture with respect to longevity and the present value of age-specific VSL. For example, if $v(t)$ is constant and $\rho(t) = exp(-\omega t)$ corresponds to exponential discounting at rate ω, then $u(t) \propto -sgn(\omega) \cdot exp(-\omega t)$, i.e., it exhibits constant absolute risk aversion with coefficient ω. If $v(t)$ is constant and $\rho(t) = t^{-\gamma}$ corresponds to hyperbolic discounting, then $u(t) \propto t^{1-\gamma}/(1-\gamma)$, i.e., it exhibits constant relative risk aversion with coefficient γ. Risk neutrality with respect to longevity implies that $v(t)$ is inversely proportional to the discount factor; e.g., with exponential discounting at rate ω, VSL ($-v(t)$) increases at rate ω. Risk-seeking preferences with respect to longevity imply that VSL increases more rapidly than the discount factor $\rho(t)$ decreases.[5]

Empirical evidence on how individuals' VSL varies with age is mixed (Hammitt 2007). Standard life-cycle models and some empirical studies suggest VSL rises then falls with age, though the age at which it is maximized and how sharply it rises and falls are uncertain (e.g., Shepard and Zeckhauser 1984, Aldy and Viscusi 2007, 2008, Krupnick 2007). Hence it is uncertain *a priori* what preference ordering over alternative perturbations to her survival curve an individual may hold. Moreover, the average rate of substitution of wealth for life-expectancy gain or for mortality-risk reduction within a period is not constant but is a decreasing function of the magnitude of the life-expectancy gain or risk reduction (Hammitt 2013, 2020).

3. Prior estimates of VSL and VSLY

VSL has been estimated in hundreds of studies in many countries. Many of these are hedonic-wage studies based on the relationship between workers' wages and occupational fatality risk (see Viscusi 2015 and Viscusi and Masterman 2017 for recent summaries). In simple terms, a model is estimated that describes how the wage and occupational fatality risk covary among the set of jobs available to a worker, which depends on her education, work experience, and other factors. On the assumption the worker is aware of the relationship between wages and risk and chooses the job she prefers most among those available, one can conclude that her VSL is approximately equal to the slope of the function relating wage to occupational fatality risk at her current job. Specifically, the worker prefers the job she holds to riskier, higher-wage alternatives, so her VSL is larger than the ratio of the incremental income to the incremental risk associated with those jobs; similarly, she prefers the

[5] Note that dynamic consistency implies $u(t)$ exhibits constant absolute risk aversion (including positive, negative, or zero risk aversion).

job she holds to safer, lower-wage alternatives, so her VSL is smaller than the ratio of the incremental lost income to the incremental risk reduction associated with those jobs.

Many other studies of VSL use stated-preference methods, often subdivided between contingent-valuation and choice-experiment studies (see Lindhjem et al. 2011 and Masterman and Viscusi 2018 for recent meta-analyses). Stated-preference studies describe a hypothetical decision and ask survey respondents which alternative they would choose. These are usually framed as willingness to pay (WTP) for a risk reduction, i.e., the compensating variation for the risk reduction. (A modest number of studies have elicited willingness to accept compensation to forgo a risk reduction, i.e., the equivalent variation. These estimates are generally larger than WTP for reasons that are not well understood; see Tunçel and Hammitt 2014 for a recent meta-analysis.) In contingent valuation, the choice is binary, often between the status quo and an alternative; in choice experiments, there are usually three or more alternatives (often one is the status quo). Some contingent-valuation studies ask the respondent her maximum WTP as an open-ended question or as a choice among several proposed amounts or ranges. To estimate VSL, the alternatives differ in mortality risk and their effect on individual wealth (through a product price, change in taxes, or other payment vehicle). Again, each respondent is assumed to prefer the alternative she chooses to the others offered, from which one can infer a bound on her VSL. The decision context used in stated-preference studies varies widely, including choices among foods and drinking-water sources, transportation alternatives, medicines, and unspecified mechanisms.

One of the major concerns with stated-preference studies is whether responses can be interpreted as consistent with a respondent's reflective (thoughtful, informed) preferences; because the respondent faces no significant consequence from her choice, there is limited incentive to reflect carefully or even to report honestly. One criterion often recommended for evaluating stated-preference studies is whether responses are sensitive to scope, i.e., consistent with the hypothesis that WTP is an increasing function of the magnitude of the good that is valued. For reductions in mortality risk, standard models of VSL imply that WTP should be close to but less than proportionate to the magnitude of the risk reduction (Hammitt and Graham 1999, Corso et al. 2001, Alolayan et al. 2017, Hammitt and Herreira 2018, Hammitt et al. 2019, Hammitt 2020).

Estimates of VSLY are much less common than of VSL. VSLY is often defined as the rate of substitution between wealth and the expected discounted present value of longevity. For hedonic-wage studies, VSLY has been estimated by substituting for current risk an estimate of the product of occupational risk and the expected present value of longevity (often approximated as discounted life expectancy, Jones-Lee et al. 2015); the relationship between wage and risk can be interpreted as VSLY (Moore and Viscusi 1988, Aldy and Viscusi 2007, 2008). Similarly, VSLY is often estimated by dividing an estimate of population-mean VSL by population-mean life expectancy (e.g., Hirth et al. 2000). Mason et al. (2008) estimate VSLY by dividing VSL by life expectancy and, alternatively, by interpreting the estimated change in VSL with age as a measure of the value of the associated decrease in life expectancy. Their second approach implies negative VSLY for ages over which estimated VSL increases with age.

A few stated-preference studies have estimated VSLY directly, by asking respondents to choose between alternatives described by their effects on life expectancy and on wealth. Perhaps the earliest examples are Johannesson and Johansson (1996, 1997) who used binary-choice questions to elicit current WTP by adults for an increase in life expectancy at age 75 from 10 to 11 years. The estimated VSLY is $1500 or less, a surprisingly small amount. Morris and Hammitt (2001) elicited current WTP for a pneumonia vaccine from four independent subsamples. The vaccine was describing as being taken at age 60 or 70 years and the benefit was described either as a reduction in the average annual mortality risk after that age (from 4.8 to 4.6 percent and from 7.0 to 6.8 percent, respectively) or as an increase in life expectancy at that age (from 21 years to 21 years 11 months, and from 14 years to 14 years 5 months, respectively). Morris and Hammitt found no significant difference in WTP for the vaccine offered at 60 or 70 years when the benefit was described as a risk reduction but that WTP was about 1.6 times larger when the benefit was described as an increase in life expectancy; this suggests that life expectancy might be a more easily understood summary of the effect of mortality risk reduction. A weakness of this study is that the change in annual risk was the same for vaccines beginning at ages 60 and 70; the benefit of the earlier intervention is that one benefits from the risk reduction for ages 60 to 69, as well as for ages 70 and older.

Desaigues et al. (2011) describe a stated-preference study designed to value the change in life expectancy associated with a reduction in air pollution. The scenario that was presented to respondents was complex and intended to be realistic: ambient air pollution would fall linearly for 20 years (by either 1.5 or 3.0 percent of the initial level per year, yielding a 30 or 60 percent total reduction) and would remain constant thereafter. The intervention would produce an increase in life expectancy of 3 or 6 months. The effect of air pollution was characterized as affecting “ability to survive,” a qualitative concept illustrated by a graph that declines with age, reaching zero at death. A reduction in air pollution shifts the ability-to-survive function to older ages, which was intended to convey to respondents that there is an improvement in health at all ages. The survey elicited monthly WTP for the rest of one’s life and was administered in nine major European cities. Mean WTP for the subsample that valued the 6 month life-expectancy increase was about 1.3 times as large as mean WTP for the subsample that valued the 3 month life-expectancy increase; using the (larger) estimates from the 3 month increase, mean VSLY is estimated as 41,000 euros in the 15 original European Union member states plus Switzerland and 33,000 euros in three newer member states.

Cameron and DeShazo (2013) report an elaborate stated-preference survey, fielded by internet in the US and Canada. They elicited WTP to reduce the risk of suffering each of a broad set of "illness profiles" defined by the allocation of life years over four categories: before, during, and after illness, and lost to premature death. The values of years in each category can depend on the allocation across categories as well as on the characteristics of the illness (e.g., disease, treatment) and of the individual (e.g., age, income). Their empirical model represents individual utility as a function of the logarithm of income and of life years in each category, which implies that average VSLY is a decreasing function of life years saved. Cameron and DeShazo report estimates of the marginal rate of substitution of income for reduction in the risk of illustrative illness profiles. For example, for a 45 year old with average income, the values of reducing by 1 per million the probabilities of immediate death, 1 year of illness then death, and 5 years of illness then death, are $6.74, $8.09, and $9.09, respectively (their Table 2, 5 percent discount rate; these can be scaled as VSL by multiplying by one million). The average VSLY for each of these scenarios can be derived with knowledge of the life expectancy absent illness; for the immediate death scenario, it is roughly $200,000 (assuming life expectancy of 35 years at age 45 and no discounting, i.e., $6.74 million/35 years). The reported values imply that living 1 or 5 years with fatal illness then dying is worse than dying at the age when the illness would begin; i.e., the average value per year lived with illness is less than the average value per life year lost to death (for these illness profiles and individual characteristics).

4. Survey design and administration

The survey was designed to elicit individuals' preferences between, and WTP for, alternative perturbations to their hazard functions. The survey instrument and administration are described in the following subsections.

4.1. Survey instrument

The survey instrument consists of five sections. Respondents were informed that the survey would begin with some explanatory material about life expectancy followed by questions about their preferences for increasing life expectancy by decreasing mortality risk at different ages. They were assured that responses depend on their preferences and that there are no right or wrong answers.

The building blocks of the survey are baseline mortality hazard functions (for each gender and 10-year age group) and nine perturbations to each baseline hazard. The perturbations differ in the time path of risk reduction and life-expectancy gain. Baseline hazard is represented as the average annual mortality risk for each 10-year period. The three perturbations are “transient” (T) that decreases mortality risk for the first decade, “additive” (A) that decreases mortality risk by subtracting a constant from the risk in every decade, and “proportional” (P) that decreases mortality risk by multiplying the risk in each decade by a constant. In all cases, the perturbation begins at age *m*, where *m* is the smallest multiple of 10 greater than the individual’s current age (e.g., *m* = 30 for an individual aged 20-29 years).[6]

The hazard functions are:

$$\text{Transient:} \quad h_t(t) = h_0(t) - c \qquad m \le t < 10 \qquad (7a)$$
$$= h_0(t) \qquad t \ge 10$$

$$\text{Additive:} \quad h_a(t) = h_0(t) - a \qquad t \ge \mathrm{m} \qquad (7b)$$

$$\text{Proportional:} \quad h_p(t) = h_0(t)\,(1 - p) \qquad t \ge \mathrm{m} \qquad (7c)$$

where $h_0(t)$ is the baseline hazard function, and *c*, *a*, and *p* are positive numbers that depend on age and gender.

These patterns are representative of the effects of many types of interventions: those that decrease risk of cardiovascular disease, many types of cancer, and other diseases of old age yield an increasing hazard reduction with age (like the proportional perturbation); those that reduce external hazards such as increasing transportation safety or protection against fires may have a roughly constant effect across ages (like the additive perturbation), and those that affect only a particular time interval (e.g., a temporary occupation, a disease outbreak, the lifetime of a motor-vehicle) have only a transient effect.

There are three levels of life-expectancy gain. The perturbations are standardized across age groups and gender by applying a common factor for the proportional pattern *p* = 0.15, 0.38, or 0.60 percent. These factors produce life-expectancy gains that are larger for younger age groups (who experience the risk reduction for more decades) and for men (who have a larger baseline hazard). The life-expectancy gains for the three risk reductions were presented as whole numbers of days and are between 6 and 8, 15 and 20, and 24 and 32 days, respectively. In the following, we label the possible life-expectancy gains by their averages across age groups and gender, about 7, 18, and 28 days, respectively.[7] The transient and additive perturbations are constructed for each age and gender group to produce the same life-expectancy gain as the proportional perturbation for that group. The

[6] For all perturbations, the hazard equals baseline hazard for ages less than *m*.
[7] The exact figures that are used to calculate VSLY are 6.9, 17.7, and 28.0 days.

risk reductions and life-expectancy gains are small enough to be realistic and relevant for policy evaluation: for example, the proportional perturbation could be achieved by decreasing ambient exposure to fine particulate air pollution ($PM_{2.5}$) by 1 μg/m^3 or less.[8]

The introductory screen described life expectancy as the number of future years one can expect to live. It illustrated annual mortality risk and survival probability (averaged by decade) and life expectancy for an average individual of the respondent's age and gender (Figure 1 provides an example). It noted that annual mortality risk is very low when young and increases with age, and that decreasing mortality risk at any age increases life expectancy and the chance of living at older ages.

The second section stated that different programs could reduce mortality risk more or less at different ages: "For example, making food or transportation safer might decrease the risk of dying in a year by about the same amount when one is young or old. As another example, better treatments for heart attacks would reduce the risk of dying in a year more at older ages, because people rarely suffer heart attacks when they are young." It then introduced three programs (labeled X, Y, and Z) corresponding to the transient, additive, and proportional perturbations. Each program increased life expectancy by the same amount (randomly selected for each individual from the three possible values). The programs were accompanied by graphics illustrating the baseline hazard, perturbed hazard, change in annual mortality risk, baseline and increase in life expectancy, and simple arguments in favor of and against the program. Figure 2 provides an example.

The third section presented respondents with all three pairwise choices among the programs. Respondents were asked to assume they could benefit from each program at no cost and to state which they preferred, or if they were indifferent between them (Figure 3 provides an example). The life-expectancy gain for each respondent was identical between programs and randomly selected from the three possible levels. There are 27 possible response patterns (including indifference), of which 13 are transitive. Respondents who are indifferent among the three programs are risk-neutral with respect to longevity; those who strictly prefer transient to additive to proportional are risk-averse and those who strictly prefer proportional to additive to transient are risk-seeking. The order of pairs and the ordering between programs in each pair were randomized with the constraint that a respondent who always chose the first program (or always chose the second) in each pair would reveal an intransitive ranking.

[8] Relative mortality risk is estimated as 1.064 per 10 μg/m^3 (Héroux et al. 2015). US average concentration decreased from about 13 to 8 μg/m^3 between 2000 and 2020 (https://www.epa.gov/air-trends/particulate-matter-pm25-trends).

The fourth section elicited willingness to pay (WTP) for each of three programs differing in both perturbation (T, A, P) and life-expectancy gain (approximately 7, 18, or 28 days). Each respondent answered three valuation questions: one about each perturbation and each gain in life expectancy (the pairing of these attributes and the order of presentation were randomized across respondents). WTP was elicited using double-bounded binary-choice questions (Hanemann et al. 1991). As a validity test, half the respondents were asked about WTP per year and half about WTP per month; in both cases payments would be required for the next 10 years. Response options were "Pay for the program every year [month] and increase my life expectancy" and "Not pay for the program, and not increase my life expectancy." The initial bid was randomly selected from a set ranging between $20 and $8000 per year ($2 and $800 per month), drawn without replacement so each respondent was presented with different bids for each perturbation/life-expectancy gain. The follow-up bid was half as large if the respondent rejected the initial bid and twice as large if she accepted it. Respondents who rejected both initial and follow-up bids were asked if they would accept the program if it were free. Respondents who reject the program if it were free can be interpreted as violating the assumption that the value of a life-expectancy gain is non-negative or as providing a protest response indicating rejection of some aspect of the survey (analogous to "protest zeros" in an open-ended WTP question).

The final section asked respondents about their health and life satisfaction, both at present and anticipated at age 80, and about their perceived chance of living to age 80 compared with others of their age and gender. Health was elicited using both the standard categorical scale (excellent, very good, good, fair, poor) and a visual analog scale (between full health = 100 and as bad as dead = 0). Life satisfaction used an integer scale (from 0 to 10) and chance of living to 80 a five-point categorical scale (1 = much larger, 2 = a little larger, 3 = about the same, 4 = a little smaller, 5 = much smaller).

A perennial concern with stated-preference studies is that results can be invalid because respondents face little incentive to carefully consider their preferences or to respond truthfully. Several authors have presented evidence that beginning the survey by inviting respondents to take a solemn oath stating that they will respond honestly can produce results that are more plausibly interpreted as valid measures of WTP (e.g., Carlsson et al. 2013, Jacquemet et al. 2013, 2017, de-Magistris and Pascucci 2014) or are more consistent with results from incentivized experiments (Jacquemet et al. 2019). To test this effect, after being welcomed to the survey and told the compensation they would receive for completing it, half the respondents (randomly selected) were presented with a screen presenting the oath "I promise that throughout this entire survey I will tell the truth and always provide honest answers," to which they were required to respond either "yes" or "no."

4.2. Survey administration

The survey was administered to the AmeriSpeak online panel maintained by NORC. AmeriSpeak is widely used and nationally representative. Panel members are recruited using area probability and address-based sampling; the panel covers 97 percent of US residents with known, non-zero sampling probability.[9] Panel members are compensated for completing surveys (with reward points worth $3 for this study).

The sample was restricted to individuals aged 20 to 69 years and stratified to yield equal numbers of men and women in each 10-year age group (20-29, …, 60-69 years). It was fielded in several waves. A pre-test (February 27 - March 4, 2020) yielded 169 completions (response rate = 38.5 percent). No changes were made to the survey instrument after the pre-test and these respondents are included in the full sample. The first wave (March 19 - March 24, 2020) yielded 233 completions (response rate = 33.4 percent). Because of concern that responses might be influenced by the rapidly developing covid-19 pandemic, data collection was paused and ultimately resumed in early 2021 (February 11 - April 6), yielding 664 completions (response rate = 43.7 percent).[10]

5. Results

In the following subsections, we report results of the pairwise choices between time paths of risk reduction, regression models that describe WTP as a function of the time path of risk reduction and increase in life expectancy, and the corresponding VSLY. In total, 1052 respondents completed the survey.

5.1. Pairwise choices and risk posture

Respondents were presented with three pairwise choices between the transient, additive, and proportional perturbations with a common increase in life expectancy (about 7, 18, or 28 days, randomly selected for each respondent). Based on their responses, we categorize respondents into subgroups reflecting their expressed preferences over alternative time paths of risk reduction. As summarized in Table 1, 88 percent of respondents made choices consistent with a transitive preference ordering. The probability of such a large fraction exhibiting transitive preferences if

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[9] For details, see "Technical Overview of the AmeriSpeak® Panel: NORC's Probability-Based Household Panel" updated January 26, 2021, available at https://amerispeak.norc.org/us/en/amerispeak/research.html.

[10] The target sample size of 1000 was reached on March 5 but fielding continued until all the age/gender cells were complete.

responses to the pairwise questions were random is infinitesimal: the z statistic exceeds 25. Respondents who are less certain about their preferences or less attentive to the survey questions are more likely to be classified into the intransitive or risk-neutral subgroups. As noted above, pairwise choices were presented so that a respondent who always chose the first (or always chose the second) alternative would reveal an intransitive preference order. A respondent who reported indifference as an expression of uncertainty would be classified as risk neutral.

Of the full sample, 13 percent reported indifference among the three programs, consistent with risk neutrality with respect to longevity; 20 percent preferred proportional to additive to transient, consistent with risk seekingness, and 12 percent preferred transient to additive to proportional, consistent with risk aversion. Thirteen percent preferred additive to proportional to transient (henceforth "A>P>T"), which is not consistent with any global risk posture, and 27 percent made pairwise choices consistent with other transitive orderings. These frequencies are similar to the results obtained by Hammitt and Tunçel (2015) in an online survey of 1024 adults from a panel representative of the French general population and by Nielsen et al. (2010) in an in-person survey of 129 roughly 40-year-old Newcastle-area (UK) residents. Compared with the French sample, we have fewer respondents who exhibit risk neutrality (13 v. 23 percent) and more who exhibit other transitive patterns (27 v. 19 percent).[11]

Table 1 reports descriptive statistics for the full sample and for six subgroups defined by preference ordering. Descriptive statistics are similar across subgroups with a few exceptions. Respondents who are married or living with a partner are disproportionately represented in the risk-averse group and those with children younger than 18 are over-represented in the risk-averse and A>P>T groups while under-represented in the other-transitive group. College graduates are under-represented in the intransitive group and somewhat over-represented in the risk-seeking, risk-averse, and A>P>T groups. Self-employed respondents are over-represented in the risk-neutral group. Retirees are over-represented in the risk-neutral group and under-represented in the risk-seeking and A>P>T groups. On average, respondents in the risk-neutral and intransitive subgroups are older and those in the risk-seeking subgroup are younger. There are no large differences in self-rated health, life satisfaction, or perceived life expectancy among groups: Current health averages about 80 (out of 100) while expected health at age 80 is much smaller, about 60. Life satisfaction is expected to decrease by a smaller fraction, from a current average of 7.5 (out of 10) to 7.0. Perceived life

[11] Frequencies (percent) of inferred preference ordering from Hammitt and Tunçel (2015) and from Nielsen et al. (2010), respectively, are: risk-neutral 23, 6; risk-seeking 16, 23; risk-averse 14, 22; A>P>T 13, 17; other transitive 19, 23; intransitive 15, 9.

expectancy is slightly greater than for others of the same age and gender, averaging 2.7 (where 1 = much larger and 3 = about the same).

5.2. Estimated WTP for risk reduction

We estimate regression models to describe WTP in the full sample and independently in each large subgroup (as defined by responses to the pairwise choices). In the subgroups, we estimate:

$$\log(w_{ijk}) = \alpha + \sum_{j=1}^{2} \beta_j p_j + \sum_{k=1}^{2} \gamma_k l_k + X_i \theta + \varepsilon_{ijk} \tag{8}$$

where w_{ijk} is individual *i*'s WTP for perturbation p_j with life-expectancy gain l_k, X_i is a vector of individual and survey characteristics, ε_{ijk} is an independently, identically, and normally distributed error term, and α, β_j, γ_k, and θ are coefficients to be estimated. In the full sample, we add indicator variables for the subgroups of interest (risk-seeking, risk-averse, A>P>T, other-transitive, and intransitive) together with interactions between these and indicator variables for the transient and proportional perturbations. The omitted categories for perturbation, life-expectancy gain, and subgroup (for the full-sample model) are the intermediate categories: additive perturbation, life-expectancy gain = 18 days, and risk-neutral. Taking log(WTP) as the dependent variable is motivated by the assumptions that WTP should be close to proportional to the gain in life expectancy, that the monetary value of one perturbation over another should increase with the magnitude of the risk reduction, and that the error term is likely to be proportional rather than additive. Because WTP is interval-censored,[12] equation (8) is estimated using maximum likelihood (Alberini 1995).

Each respondent contributes three observations corresponding to her WTP for each of three combinations of perturbation and life-expectancy gain. Recall that each WTP question presented to a respondent used a different perturbation type, life-expectancy gain, and initial bid. Hence estimated effects of perturbation or life expectancy gain are identified by differences in WTP between, rather than within, respondents. This choice helps protect against bias due to arbitrary coherence of responses (Ariely et al. 2003).

We exclude responses to 476 valuation questions (15 percent) for which the respondent rejected the risk reduction when it was free. As noted above, these responses are inconsistent with the assumption that the value of risk reduction is non-negative and suggest the respondent rejected some aspect of the valuation question (or was inattentive).

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[12] WTP is assumed to be greater than any bid the respondent accepted (greater than zero if she rejected both bids) and less than any bid the respondent rejected (unbounded above if she accepted both bids).

Estimates for the basic specification excluding individual and survey characteristics X_i are presented in the first column of Table 2A for the full sample and in Table 2B for the independent estimates by subgroup. The estimated coefficients on life-expectancy gain are statistically significantly different from zero in the full sample (Table 2A), which implies that WTP is significantly related to life-expectancy gain. If WTP were proportional to life-expectancy gain, the coefficients for the 7- and 28-day gains would be -0.942 and 0.459, respectively. For the 7-day gain, we can reject the hypothesis that the estimated coefficient equals this value at the 1 percent level (one-sided test) but for the 28-day gain we cannot reject the hypothesis of proportionality. The estimated coefficients on life-expectancy gain differ between subsamples (Table 2B). For the 7-day gain, only the coefficients for the risk-seeking and intransitive subgroups are significantly different from zero and we can reject the hypothesis that WTP is proportional to life-expectancy gain for the risk-averse, A>P>T, and other-transitive subgroups. For the 28-day gain, only the coefficients for the A>P>T and other-transitive groups are significant and we cannot reject the hypothesis that WTP is proportional to life-expectancy gain for any group.

Evidence about the consistency of WTP for the different perturbations with the preference ordering implied by the pairwise choices is mixed. As shown in Table 2B, for the risk-seeking and A>P>T subgroups, WTP for the transient (least-preferred) program is significantly and substantially smaller than for the other programs. WTP for the proportional program is not significantly different from WTP for the additive program, although the signs of the estimated coefficients are consistent with each group's preference order (positive for risk-seeking, negative for A>P>T). For the risk-neutral subgroup, there are no statistically significant differences in WTP by perturbation type, consistent with risk neutrality. The intercepts suggest that average WTP is larger for the intransitive than for the other subgroups. These results are confirmed by the full-sample results (Table 2A): the estimated interactions of the transient program with the risk-seeking and A>P>T subgroup indicators are significantly negative and the indicator for the intransitive subgroup is significantly positive.

The residual standard deviation is larger for the intransitive (3.51) and risk-neutral (3.42) subgroups than for the other subgroups (3.26 and smaller) (Table 2B). This suggests greater heterogeneity or more random error in these subgroups, consistent with the possibility that respondents who are less certain of their preferences or less attentive to the survey are more likely to be classified into these subgroups.

Tables 3A and 3B report estimates for the models specified in equation (8), including respondent and survey characteristics X_i. The estimated effects of life-expectancy gain and perturbation type, and (in the full-sample model) the interactions between perturbation type and subgroup are similar to the estimates in the basic models (Tables 2A and 2B). Estimates for the full sample (Table 3A) suggest that women and whites have significantly smaller WTP than other respondents, while those with children younger than 18 years have significantly larger WTP. Subgroup estimates (Table 3B) show that the estimated effects of gender and having children differ between subgroups, while the effect of being white is consistently negative. The estimated effects of age, income, education, and marital status are not significantly different from zero in the full sample, though they are significant in some subgroups. In the full sample, WTP is significantly positively associated with anticipated future health and weakly negatively associated with current life satisfaction, but these effects are inconsistent across subgroups.

There is no significant difference in annual WTP associated with whether the valuation questions asked about monthly or annual payments. In the full sample, there is no difference in WTP between surveys conducted in Spring 2021 and those conducted one year earlier in the pandemic and no significant difference associated with respondents taking an oath to respond truthfully. The effects of both variables are significantly negative in the risk-neutral subgroup, but not in any of the other subgroups (Table 3B).

In summary, estimated WTP is significantly associated with the magnitude of the life-expectancy gain, though the effect is less than proportionate. On average, there is no significant difference in WTP by perturbation type, though there is some evidence of differences that are consistent with respondents' preference ordering over the perturbation types, e.g., the risk-seeking and A>P>T subgroups express significantly smaller WTP for the transient (least-preferred) risk reduction.

5.3. Estimated VSLY

Estimates of WTP from the regressions specified by equation (8) are sensitive to the method used for retransformation from the dependent variable, log(WTP), to dollars. Assuming the error ε_{ijk} is normally distributed implies WTP is lognormal. For any perturbation and life-expectancy gain, predicted median WTP is obtained by predicting log(WTP) using the estimated regression coefficients and exponentiating; mean WTP is obtained by multiplying the predicted median by $exp(\hat{\sigma}^2/2)$ where $\hat{\sigma}$ is the estimated residual standard deviation. The estimate of $\hat{\sigma}$ is 3.25 in the simple full-sample model (first column of Table 2A). Using this value, $exp(\hat{\sigma}^2/2) \approx 200$. Given this large difference between predicted mean and median WTP, we supplement our primary model with a linear regression identical to equation (8) but using WTP (w_{ijk}) as the dependent variable. Results for this linear model are presented in the second column of Table 2A. The estimated effects of life-expectancy gain are of the correct sign, but neither is significantly different from zero. The only statistically significant effects are that average WTP is larger for the intransitive subgroup and that WTP in the A>P>T subgroup is smaller for the transient and proportional than for the additive perturbation, consistent with the subgroup's preference ordering. Except for the interaction of intransitive subgroup with the proportional perturbation, these effects are also statistically significant in the logarithmic model (the first column of Table 2A).

To estimate VSLY, we predict mean annual WTP for each perturbation and life-expectancy gain using the full-sample regression models reported in Table 2A, multiply by 10 to produce total WTP (because respondents are told they would pay for 10 years[13]) and divide by the life-expectancy gain (measured in years).

The top panel of Table 4 presents the calculated VSLY using the logarithmic model. The first nine rows show the results for each perturbation type, life-expectancy gain, and subgroup. The following three rows show the mean values for each life-expectancy gain (averaging over perturbation type) and the following two rows show the arithmetic and geometric means of the mean values by life-expectancy gain. Within each row, the last two columns report the means (weighted by subgroup frequency) for all subgroups and excluding the intransitive subgroup. Estimates of VSLY from the intransitive subgroup seem less valid, as these respondents do not exhibit a coherent preference ordering over the alternative perturbations and exhibit greater variability in valuation (reflected by the larger residual standard deviation than for other subgroups, Table 2B). The bottom panel (last five rows) presents results calculated using the linear model. We omit the detailed results by perturbation type and report only the mean values for each life-expectancy gain and the arithmetic and geometric means of these.

Across all perturbations, life-expectancy gains, and subgroups, VSLY calculated from the logarithmic model ranges from about $200,000 to $2.5 million ($1.2 million excluding the intransitive subgroup). Because predicted WTP is less than proportional to life-expectancy gain, the calculated VSLY is smaller for larger gains in life expectancy. Averaging across perturbations, calculated VSLYs for 7, 18, and 28-day gains in life expectancy equal $833,000, $429,000, and $400,000, respectively, for the transitive subgroups.[14]

Given the less-than-proportional relationship of WTP to life-expectancy gain, the most-reasonable summary measure combining the different life-expectancy gains is the geometric mean. For the logarithmic model, this summary ranges from $387,000 to $570,000 for the transitive subgroups, to $1.5 million for the intransitive group. Combining subgroups, the geometric mean across life-

[13] If respondents discount future payments, VSLY will be smaller than we report by a factor equal to the present value of 10 equal annual payments divided by 10. For annual discount rates of 3 and 5 percent, the corresponding factors are 0.95 and 0.87, respectively. Hence discounting future payments would have only a modest effect on calculated VSLY.

[14] Predicted median VSLYs are 200 times smaller.

expectancy gains is $636,000 for the full sample and $523,000 for the transitive subgroups. The variation in VSLY by subgroup is modest; among the transitive subgroups, the geometric mean VSLY is within 10 percent of the full-sample value, except the value for the risk-neutral subgroup is about 25 percent smaller. In contrast, the value for the intransitive subgroup is almost thee times larger.

Calculated VSLY using the linear model is modestly smaller than using the logarithmic model. As shown in the bottom rows of Table 4, the mean VSLY by life-expectancy gain for the transitive subgroups ranges from $231,000 to $807,000, with a geometric mean of $393,000, about 25 percent smaller than the corresponding value from the logarithmic model ($523,000). Calculated VSLY is more sensitive to life-expectancy gain when calculated from the linear model than when calculated from the logarithmic model, because the estimated proportional effects of life-expectancy gain on WTP are smaller. Similarly, the differences in VSLY between subgroups are also smaller when calculated from the linear model; the geometric mean VSLY for each transitive subgroup is within 10 percent of the average for the transitive subgroups and the average for the intransitive subgroup is only 40 percent larger.

These estimates of VSLY are somewhat larger than conventional estimates obtained by dividing VSL by life expectancy at the average age. For example, dividing a VSL of $10 million by remaining life expectancy of 40 years yields VSLY equal to $250,000; discounting future years at 3 percent annually yields a value of more than $400,000.

6. Conclusions

The value of reducing mortality risk to an individual can be described alternatively using the concepts of VSL and VSLY. In general, these values depend on the individual's baseline mortality hazard (and life expectancy) as well as her lottery on future income, health, and other factors. Few studies have estimated VSLY directly; most estimates are derived by dividing VSL by an estimate of remaining life expectancy (or of the expected present value of discounted longevity). This study is the first, to our knowledge, to directly estimate WTP for a diverse set of risk reductions, characterized by the time path of reductions in mortality hazard and the increase in life expectancy.

We find that individuals' preferences for different time paths of risk reduction producing equal gains in life expectancy are heterogeneous. The frequency distribution of different preference orderings is broadly similar to those found in previous studies (Nielsen et al. 2010, Hammitt and Tunçel 2014). Between 12 and 20 percent of respondents made pairwise choices between alternative perturbations that are consistent with global risk neutrality, risk seekingness, risk aversion, or other preference orderings. In total, almost 90 percent of respondents made pairwise choices that are transitive. The probability of observing such a large fraction if responses to the pairwise choices were random is infinitesimal.

WTP for risk reduction is sensitive to life-expectancy gain over the range we tested (between about 7 and 28 days). In our full-sample model, the estimated coefficients on life-expectancy gain are statistically significantly different from zero and we cannot reject the hypothesis that the difference in WTP between 18 and 28-day gains is consistent with proportionality of WTP to life-expectancy gain (though we do reject this hypothesis for the difference in WTP between 7 and 18-day gains).

There is some evidence of differences in WTP for different time paths of risk reduction that is consistent with respondents' preference ordering over the perturbations. Specifically, WTP for the transient (least-preferred) perturbation is significantly smaller than for the other perturbations in the risk-seeking and A>P>T subgroups. WTP for risk reduction is larger for the intransitive than for the other subgroups, but estimates from this subgroup are less compelling because respondents may be more uncertain about their preferences or less attentive to the survey.

Our estimates of VSLY are comparable to, but somewhat larger than conventional estimates obtained by dividing estimated VSL by average life expectancy. Estimates based on predicted mean WTP from the logarithmic model range between $387,000 and $570,000 in the transitive subgroups, with an overall mean of $523,000. Using the linear regression model, the estimates range between $364,000 and $430,000, with an overall mean of $393,000.

Because the value of any change to an individual's hazard function can be characterized as a VSL or VSLY, estimates of these concepts are complementary. Better understanding of the value of mortality-risk reduction and how it varies with age and other dimensions may be achieved by comparing direct estimates of the two concepts. Consistent evidence of heterogenous preferences over alternative time paths of risk reduction suggests that accurate valuation of risk reductions must recognize this heterogeneity; one size will not fit all.

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Table 1. Descriptive statistics and risk posture with respect to longevity

	All	Risk neutral	Risk seeking	Risk averse	A>P>T	Other transitive	Intransitive
Number	1052	147	214	127	144	292	128
Share (%)	100	13	20	12	13	27	12
Female	0.48	0.48	0.42	0.49	0.54	0.50	0.46
Married or with partner	0.54	0.55	0.53	0.63	0.54	0.50	0.53
Have children < 18 yrs	0.19	0.18	0.22	0.25	0.25	0.14	0.16
White	0.68	0.68	0.73	0.65	0.73	0.65	0.62
College graduate	0.41	0.37	0.49	0.50	0.50	0.37	0.25
Employed	0.58	0.48	0.66	0.58	0.63	0.59	0.52
Self-employed	0.10	0.14	0.08	0.11	0.09	0.11	0.09
Retired	0.11	0.19	0.07	0.11	0.08	0.11	0.12
Age (years)	46.7	51.7	42.4	48.0	44.4	46.1	51.0
	(14.5)	(13.0)	(14.1)	(14.7)	(13.7)	(14.7)	(14.1)
Current health (0-100)	79.4	77.9	80.5	80.8	79.3	80.1	76.9
	(15.4)	(17.7)	(14.1)	(13.3)	(16.0)	(15.2)	(16.3)
Future health (0-100)	58.3	58.4	58.9	56.3	57.7	58.3	59.6
	(22.4)	(25.2)	(22.9)	(20.8)	(20.0)	(22.2)	(22.9)
Life satisfaction (0-10)	7.5	7.2	7.5	7.4	7.5	7.6	7.4
	(1.8)	(2.2)	(1.7)	(1.7)	(1.7)	(1.8)	(2.0)
Future life satisfaction	7.0	6.7	7.3	6.8	7.2	7.1	6.8
(0-10)	(2.2)	(2.7)	(1.9)	(2.1)	(2.0)	(2.2)	(2.4)
Perceived LE (1=much	2.7	2.7	2.5	2.6	2.8	2.7	2.6
larger, 5=much smaller)	(1.0)	(1.0)	(1.0)	(1.0)	(0.9)	(1.0)	(1.1)
Income ($000/yr)	76.69	73.45	79.15	82.26	84.31	74.20	67.89
	(49.60)	(50.18)	(51.63)	(48.90)	(52.20)	(48.36)	(44.52)

Notes: Standard deviations in parentheses. Future health and future life satisfaction are expected levels at age 80. Perceived life expectancy is compared with others of same age and gender.

Table 2A. Estimated log(WTP) and WTP, basic model

Dependent variable	Log(WTP)	WTP
Transient	0.007	17.4
	(0.257)	(196.8)
Proportional	0.040	213.9
	(0.255)	(195.6)
LE 7	-0.279*	-55.8
	(0.183)	(138.5)
LE 28	0.388**	186.8
	(0.178)	(137.5)
RS	0.599	-39.7
	(0.389)	(294.4)
RA	0.503	-94.3
	(0.447)	(340.1)
A>P>T	0.666	297.3
	(0.434)	(329.1)
Other transitive	0.373	-43.8
	(0.272)	(205.0)
Intransitive	1.361***	494.3**
	(0.318)	(246.1)
Trans*RS	-0.876*	-481.4
	(0.465)	(353.8)
Prop*RS	0.005	-218.9
	(0.456)	(351.8)
Trans*RA	-0.213	-52.8
	(0.563)	(427.5)
Prop*RA	-0.457	-226.5
	(0.559)	(427.4)
Trans*A>P>T	-1.005*	-806.7**
	(0.543)	(408.3)
Prop*A>P>T	-0.654	-838.2**
	(0.532)	(406.7)
Intercept	4.347***	1655.7***
	(0.299)	(224.7)
σ	3.250***	2770.5***
	(0.065)	(44.5)
N	2676	2676

Notes: *, **, *** denote significantly different from zero at 10, 5, 1 percent, respectively. Significance tests for LE 7 and LE 28 are one-sided.

Table 2B. Estimated log(WTP) by subgroup, basic mode

	Risk neutral	Risk seeking	Risk averse	A>P>T	Other transitive	Intransitive
Transient	0.433	-0.863**	-0.100	-1.005**	-0.109	-0.112
	(0.609)	(0.372)	(0.499)	(0.464)	(0.343)	(0.568)
Proportional	0.541	0.038	-0.290	-0.575	0.079	-0.540
	(0.604)	(0.363)	(0.497)	(0.454)	(0.339)	(0.566)
LE 7	-0.752	-0.512*	0.352	0.124	-0.126	-1.022**
	(0.616)	(0.371)	(0.503)	(0.463)	(0.343)	(0.572)
LE 28	0.241	0.182	0.118	0.813**	0.741**	-0.164
	(0.600)	(0.361)	(0.463)	(0.453)	(0.335)	(0.559)
Intercept	4.203***	5.113***	4.667***	4.755***	4.571***	6.363***
	(0.553)	(0.330)	(0.465)	(0.419)	(0.316)	(0.529)
σ	3.420	3.088	3.210	3.126	3.260	3.512
	(0.203)	(0.134)	(0.178)	(0.162)	(0.121)	(0.223)
N	269	590	350	399	763	305

Notes: *, **, *** denote significantly different from zero at 10, 5, 1 percent, respectively. Significance tests for LE 7 and LE 28 are one-sided.

Table 3A. Estimated log(WTP) with covariates

Transient	0.035	Female	-0.336**
	(0.254)		(0.148)
Proportional	0.054	log(income)	0.040
	(0.252)		(0.104)
LE 7	-0.281*	College graduate	-0.156
	(0.180)		(0.157)
LE 28	0.399**	White	-0.859***
	(0.176)		(0.166)
RS	0.701*	Married, partner	-0.222
	(0.392)		(0.175)
RA	0.634	Children < 18 yrs	0.536***
	(0.446)		(0.197)
A>P>T	0.869**	Current health	0.0006
	(0.434)		(0.0011)
Other transitive	0.469*	Future health	0.0054***
	(0.275)		(0.0015)
Intransitive	1.400***	Life satisfaction	-0.305*
	(0.320)		(0.175)
Trans*RS	-0.891*	Future life sat	-0.103
	(0.459)		(0.168)
Prop*RS	0.003	Perceived LE	0.117
	(0.450)		(0.157)
Trans*RA	-0.234	Oath	-0.135
	(0.555)		(0.147)
Prop*RA	-0.511	Bid annual	0.073
	(0.552)		(0.147)
Trans*A>P>T	-1.945*	Spring 2021	-0.070
	(0.535)		(0.155)
Prop*A>P>T	-0.649	Intercept	4.746***
	(0.524)		(1.177)
Age	-0.0078	σ	3.187***
	(0.0057)		(0.064)
$(\text{Age-mean})^2$	0.0005	N	2676
	(0.0004)		

Notes: *, **, *** denote significantly different from zero at 10, 5, 1 percent, respectively. Significance tests for LE 7 and LE 28 are one-sided.

Table 3B. Estimated log(WTP) with covariates by subgroup

	Risk neutral	Risk seeking	Risk averse	A>P>T	Other transitive	Intransitive
Transient	0.391	-0.870**	-0.106	-1.070**	-0.128	-0.031
	(0.557)	(0.363)	(0.472)	(0.452)	(0.338)	(0.539)
Proportional	0.466	0.058	-0.355	-0.644	0.054	-0.429
	(0.552)	(0.353)	(0.471)	(0.442)	(0.334)	(0.539)
LE 7	-0.737*	-0.512*	0.280	0.149	-0.143	-0.949**
	(0.563)	(0.361)	(0.475)	(0.450)	(0.338)	(0.544)
LE 28	0.387	0.205	0.020	0.844**	0.740**	-0.119
	(0.548)	(0.351)	(0.467)	(0.442)	(0.330)	(0.530)
Age	-0.037*	0.0206	0.005	0.025	-0.027**	0.001
	(0.020)	(0.0135)	(0.016)	(0.015)	(0.0110	(0.020)
$(\text{Age-mean})^2$	0.0005	0.0014	-0.0005	0.0017	-0.0003	0.0026*
	(0.0013)	(0.0009)	(0.0012)	(0.0012)	(0.0008)	(0.0014)
Female	0.845*	-0.604*	-1.110***	-0.302	-0.136	-0.453
	(0.491)	(0.314)	(0.426)	(0.394)	(0.279)	(0.469)
log(income)	0.160	-0.138	0.490	-0.520	-0.175	1.074***
	(0.357)	(0.189)	(0.328)	(0.318)	(0.193)	(0.344)
College graduate	0.207	0.280	-0.414	1.086	-0.293	-1.685***
	(0.528)	(0.320)	(0.452)	(0.426)	(0.296)	(0.550)
White	-1.101*	-0.999***	-0.798*	-0.647	-0.957***	-1.580***
	(0.567)	(0.361)	(0.465)	(0.441)	(0.316)	(0.533)
Married, partner	0.213	-0.219	-0.621	-0.872	-0.671**	-1.355**
	(0.541)	(0.357)	(0.525)	(0.465)	(0.338)	(0.591)
Children < 18 yrs	-0.255	0.640	2.755***	0.847*	-0.650	0.704
	(0.622)	(0.395)	(0.513)	(0.469)	(0.436)	(0.678)
Current health	-0.0037	-0.0075	0.0065	0.0012	0.0111	0.0023
	(0.0024)	(0.0154)	(0.0200)	(0.0016)	(0.0116)	(0.0029)
Future health	0.0087***	-0.0002	0.0046	0.0041	0.0051**	-0.0024
	(0.0030)	(0.0100)	(0.0137)	(0.0113)	(0.0023)	(0.0130)
Life satisfaction	0.293	-0.420	-0.621	-0.246	-0.587	0.137
	(0.570)	(0.355)	(0.478)	(0.486)	(0.362)	(0.614)
Future life sat	-0.863	0.552	0.009	0.689	-0.316	-0.552
	(0.603)	(0.344)	(0.445)	(0.485)	(0.332)	(0.540)
Perceived LE	-0.199	0.445	-0.665	0.240	-0.229	0.803
	(0.533)	(0.334)	(0.483)	(0.440)	(0.304)	(0.528)
Oath	-1.155**	-0.428	0.108	-0.142	0.019	0.865*
	(0.498)	(0.297)	(0.446)	(0.400)	(0.281)	(0.455)
Bid annual	-0.657	-0.087	0.267	0.423	0.071	0.290
	(0.492)	(0.299)	(0.413)	(0.380)	(0.280)	(0.469)
Spring 2021	-1.641***	-0.338	0.320	0.267	0.170	0.501
	(0.481)	(0.324)	(0.417)	(0.425)	(0.297)	(0.495)
Intercept	6.339*	7.072***	-0.413	8.531**	7.625***	-5.020
	(3.830)	(2.258)	(3.572)	(3.481)	(2.224)	(3.861)
σ	2.989***	2.977***	2.956***	3.007***	3.179***	3.250***
	(0.186)	(0.131)	(0.166)	(1.570)	(0.117)	(0.206)
N	269	590	350	399	763	305

Notes: *, **, *** denote significantly different from zero at 10, 5, 1 percent, respectively. Significance tests for LE 7 and LE 28 are one-sided.

Table 4. Calculated VSLY by perturbation type and life-expectancy gain ($000)

		Risk neutral	Risk seeking	Risk averse	A>P>T	Other transitive	Intransitive	Mean	Mean (transitive)
Logarithmic model									
Transient	LE 7	612	464	818	436	888	2,386	861	665
	LE 18	315	239	421	225	458	1,229	444	342
	LE 28	294	222	392	209	426	1,145	413	319
Additive	LE 7	607	1,106	1,005	1,182	882	2,370	1,134	975
	LE 18	313	570	518	609	454	1,221	584	502
	LE 28	292	531	482	568	423	1,137	544	468
Proportional	LE 7	632	1,157	662	640	918	2,467	1,044	861
	LE 18	326	596	341	330	473	1,271	538	443
	LE 28	304	555	318	307	441	1,184	501	413
Mean by LE	LE 7	617	909	828	753	896	2,407	1,013	833
	LE 18	318	468	427	388	462	1,240	522	429
	LE 28	296	436	397	361	430	1,156	486	400
Arithmetic mean		410	604	551	501	596	1,601	674	554
Geometric mean		387	570	520	473	562	1,511	636	523
Linear model									
Mean by LE	LE 7	888	743	789	755	865	1,149	846	807
	LE 18	358	301	319	306	349	460	341	326
	LE 28	250	215	226	218	245	315	240	231
Arithmetic mean		499	420	444	426	486	641	476	455
Geometric mean		430	364	384	369	419	550	411	393

Notes: Mean by LE is arithmetic mean over perturbation type by life-expectancy gain. Arithmetic mean and geometric mean are calculated from the three means by LE. Last two columns are means of corresponding rows (weighted by subgroup frequency) except arithmetic mean and geometric mean are calculated from the three means by LE (as for the subgroups).

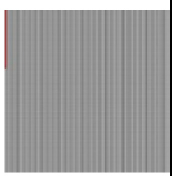
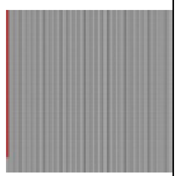
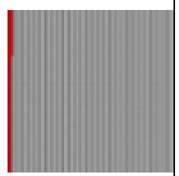
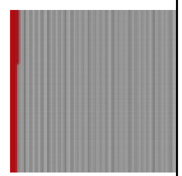
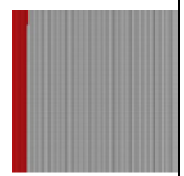
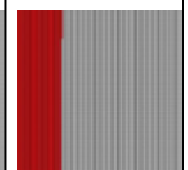
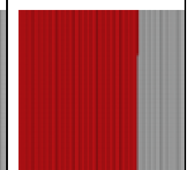
Future Decades	30 – 39 Years	40 – 49 Years	50 – 59 Years	60 – 69 Years	70 – 79 Years	80 – 89 Years	90 – 99 Years	Life expectancy
Mortality risk (out of 100 000)	585 	1 145 	2 643 	5 615 	12 153 	30 239 	68 075 	**54 years**

Figure 1. Introduction to hazard function and life expectancy (woman aged 20-29 years).

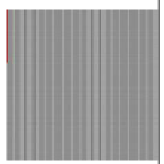
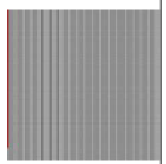
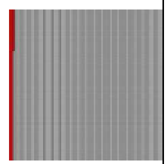
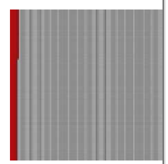
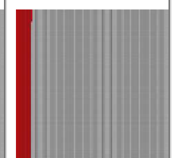
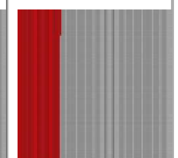
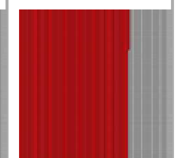
	Future Decades	30 – 39 Years	40 – 49 Years	50 – 59 Years	60 – 69 Years	70 – 79 Years	80 – 89 Years	90 – 99 Years	Life expectancy
	Mortality risk without the reduction (out of 100,000)	585 	1 145 	2 643 	5 615 	12 153 	30 239 	68 075 	**54 years**
Program X	**Mortality risk reduction (out of 100,000)**	**- 35**	**0**	**0**	**0**	**0**	**0**	**0**	**54 years + 7 days**
	New mortality risk with the reduction	**550**	**1 145**	**2 643**	**5 615**	**12 153**	**30 239**	**68 075**	

Figure 2. Introduction to program X (transient) (woman aged 20-29 years).

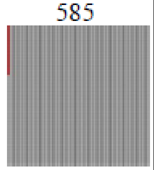
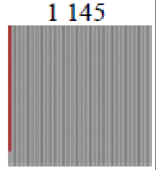
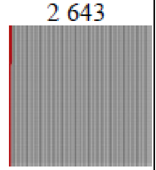
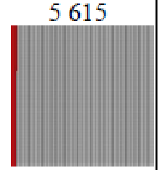
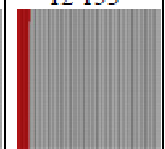
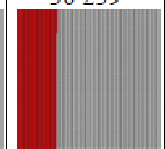
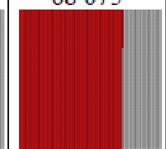
	Future Decades	30 – 39 Years	40 – 49 Years	50 – 59 Years	60 – 69 Years	70 – 79 Years	80 – 89 Years	90 – 99 Years	
	Mortality risk without the reduction (out of 100,000)	585 	1 145 	2 643 	5 615 	12 153 	30 239 	68 075 	54 years
Program Y	Mortality risk reduction (out of 100,000)	- 10	- 10	- 10	- 10	- 10	- 10	- 10	54 years + 7 days
	New mortality risk with the reduction	575	1 135	2 633	5 605	12 143	30 229	68 065	
Program Z	Mortality risk reduction (out of 100,000)	- 1	- 2	- 4	- 8	- 18	- 45	- 102	54 years + 7 days
	New mortality risk with the reduction	584	1 143	2 639	5 607	12 135	30 194	67 973	

Figure 3. Choice between programs Y (additive) and Z (proportional) (woman aged 20-29 years).