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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 with respondents' stated preferences over the alternative time paths. Estimated value per statistical life year (VSLY) can differ by time path and averages about $500,000, roughly consistent with conventional estimates obtained by dividing estimated value per statistical life by discounted life expectancy.

JEL: D61, I18, Q51

Keywords: value per statistical life, value per statistical life year, mortality risk, stated preference

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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 individuals' decisions 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 has also been investigated (e.g., Krupnick et al. 2002, Alberini et al. 2004, Hammitt and Liu 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 gains 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, Nielsen 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 sample 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.

We extend this prior research by eliciting respondents' willingness to pay (WTP) for alternative time paths of risk reduction that produce different gains in life expectancy. Our primary objective is to directly estimate VSLY (defined as WTP divided by the gain in life expectancy) and to determine how it depends on the time path of risk reduction. We evaluate the validity of these estimates by testing whether WTP is significantly associated with, and proportional to, the increase in life expectancy, and also by whether WTP for a life-expectancy gain is sensitive to whether payment is framed as occurring monthly or annually. We confirm the previous finding that respondents have transitive but heterogeneous preferences across time paths of risk reduction and extend it by testing whether respondents who prefer one time path to another in a pairwise choice express a larger WTP for the preferred time path. We note that WTP for a risk reduction that continues over many years (as elicited here) is more relevant than WTP for a short-term risk reduction (as conventionally estimated) to evaluation of environmental, health, and safety regulations and other policies that affect mortality risk, which often continue for many years.

Our data come from a stated-preference study administered online to a representative sample of about 1000 US residents aged 20 to 69 years. Respondents made pairwise choices between transient, additive, and proportional risk reductions all providing the same increase in life expectancy.[1] Similar to prior work, we find that most respondents (88 percent) report transitive and diverse preference orderings; 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. We find some evidence that differences in WTP for alternative perturbations of individuals' hazard functions are consistent with respondents' direct choices over the perturbations, but differences in WTP are imprecisely estimated. Estimated 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[2]

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

$$s(t) = 1 - \int_0^t f(x)dx \tag{1}$$

and

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

Life expectancy (remaining at *t* = 0) is the expected number of future life-years,

[1] The common life-expectancy gain ranged between about 7 and 28 days. The two prior studies used larger increases in life expectancy; about six months in Nielsen et al. (2010) and 30 to 37 days in Hammitt and Tunçel (2015).
[2] Much of this section is based on Hammitt and Tunçel (2015).

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

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. Under the standard lifecycle model, lifetime utility $u(t) = \int_0^t u'(x)dx$, where $u'(x)$ is the present value at age 0 of the utility of living at age *x,* which may depend on time preference, consumption, health, and other factors (prime denotes first derivative). We assume the individual maximizes expected lifetime utility by optimally allocating consumption over time subject to her budget constraint and the terms on which she can save or borrow and that *V* is the corresponding optimal value.

If $u(t)$ is linear, the individual is risk-neutral with respect to longevity and indifferent among all survival functions with the same 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).

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.[3] 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 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, an individual who discounts future utility might count future life years as less valuable. 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). 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

[3] 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.

investments and the terms on which she can use annuities or other tools to manage the risk of outliving her savings (Drèze 1962, Shepard and Zeckhauser 1984, 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). Pliskin et al. (1980) asked 10 members of a Harvard health-faculty group about their certainty equivalents for a lottery with equal chances of living 5 or 15 years from the present; four were risk-seeking, four were risk-neutral, and two were 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 165 respondents also chose the more risky lottery over its expected value (4 years).[4] Reed et al. (2021) found 51 percent of 200 respondents preferred a binary lottery with a 0.1 probability of 10 year survival and complementary chance of 1.1 year survival to its expected value of 2 years; 36 percent chose the fixed outcome and 14 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. The value of a small risk reduction at age t equals the product of the risk reduction and the age-specific VSL. Hence for small perturbations WTP (w) can be approximated by the expected present value of the age-dependent product of the change in hazard and the individual's VSL, i.e.,

[4] 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.

$$w = -\int_0^\infty \Delta h(t)v(t)s(t)\rho(t)dt = -\int_0^\infty \Delta f(t)v(t)\rho(t)dt, \quad (5)$$

where $\Delta h(t)$ is the change in hazard, $\Delta f(t)$ is the change in marginal probability of death, $v(t)$ is the individual's age-dependent VSL (her value of extending life beyond t), and $\rho(t)$ is her consumption-discount factor (normalized so that $\rho(0) = 1$) (Rosen 1988, Johannesson et al. 1997).[5]

Expression (5) shows that the value of a change in hazard function to an individual depends on the initial hazard (and corresponding survival function), the time path of the perturbation, the time path of her VSL (which may depend on health, income, initial hazard, 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 in 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). Moreover, for any time path of risk reduction, WTP is proportional to the increase in life expectancy (for small changes). This follows because WTP for a change in hazard at age t is proportional to the magnitude of the change; hence multiplying the change in hazard function $\Delta h(t)$ for all $t > 0$ by a constant increases both WTP (w) and the gain in life expectancy by the same factor.

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

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

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 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 different, however. The weighting factor $v(t)\ \rho(t)$ 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. Note that if $\Delta h(t) \leq 0$ for all $t > 0$, both WTP (w) and the change in expected utility (ΔV) are greater than or equal to zero. This implies that utility (which is unique up to a positive affine transformation) can be normalized so that $u(t) = -\ v(t)\ \rho(t)$ and $u(t) < 0$ for all t less than or equal to some maximum possible age M.

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

The proportionality between $u(t)$ and $-v(t)\,\rho(t)$ 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[6] and $\rho(t) = t^{-\gamma}$ ($\gamma \geq -1$) corresponds to hyperbolic discounting, then the utility function can be expressed as $u(t) = (M^{-\gamma} - t^{-\gamma})/\gamma$, i.e., it exhibits constant relative risk aversion with coefficient $\gamma + 1$.[7] Note that when $\gamma = -1$ (risk neutrality), $u(t) = t - M$ is linear. If $v(t)$ is constant and $\rho(t) = exp(-\omega t)$ corresponds to exponential discounting at rate $\omega > 0$, then $u(t) = -exp(-\omega t)$, i.e., it exhibits constant absolute risk aversion with coefficient ω. If the discount rate $\omega < 0$, $u(t) = exp(-\omega t) - exp(-\omega M)$ is risk seeking. Hence an individual who discounts future life years at a constant rate exhibits constant absolute risk aversion with respect to longevity.[8] Conversely, with exponential discounting at rate ω, risk neutrality with respect to longevity implies $v(t) \propto (M - t)e^{\omega t}$, which rises then falls with age, consistent with lifecycle models and empirical results (e.g., Shepard Zeckhauser 1984, Ng 1992, Aldy and Viscusi 2008). If $u(t) = -exp(-\alpha t)$ exhibits constant absolute risk aversion with coefficient α, exponential discounting at rate ω implies $v(t) \propto e^{(\omega-\alpha)t}$. VSL is independent of age if the discount rate and coefficient of risk aversion are equal; it decreases or increases exponentially with age if the coefficient of risk aversion is larger or smaller than the discount rate, respectively.[9]

Evidence on how individuals' VSL changes with age is mixed (Hammitt 2007). Standard lifecycle 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, Rosen 1988, Ng 1992, Aldy and Viscusi 2007, 2008, Krupnick 2007). Hence it is uncertain *a priori* what preference ordering over alternative perturbations to her hazard function an individual may hold. Moreover, the average rate of substitution of wealth for a life-expectancy gain or a 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 estimates are from 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

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[6] For notational simplicity, choose units so $v(t) = 1$ for all $t > 0$.
[7] If $\gamma = 0$, $\rho(t) = \log(t)$ and $u(t) = \log(t) - \log(M)$.
[8] An individual who discounts future years at rate ω will be indifferent among all changes to her hazard function that yield the same change in the expected present value of future life years (Jones-Lee et al. 2015).
[9] Note that dynamic consistency implies $u(t)$ exhibits constant absolute risk aversion (including positive, negative, or zero risk aversion).

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 assumptions that 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 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 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. 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 one or more bounds 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 thoughtful, informed preferences; because the respondent faces no significant consequence from her choice, there is limited incentive to reflect carefully or 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). As described above, equation (5) shows that, for any time path of risk reduction, WTP should be nearly proportional to the increase in life expectancy. WTP may differ, however, for different time paths of risk reduction even if they produce the same gain in life expectancy.

Estimates of VSLY are much less common than of VSL. Empirical studies often assume that individuals discount future life years and hence define VSLY as the rate of substitution of wealth for the expected present value of future longevity (which is often approximated as discounted remaining life expectancy, as emphasized by Jones-Lee et al. 2015). Perhaps the simplest approach to estimating VSLY is to divide the estimated average VSL by the average (discounted) life expectancy for a population (e.g., Hirth et al. 2000).[10] Mason et al. (2008) applied this approach and, as an alternative, interpreted the estimated change in VSL with age as a measure of the value of the associated decrease in life expectancy (which they note implies, problematically, that VSLY is negative for ages at which estimated VSL increases with age).

Using a hedonic-wage approach, Moore and Viscusi (1988) estimated average VSLY (and the rate at which future life years are discounted) by regressing wage on the expected loss of discounted longevity (i.e., on the product of occupational fatality risk and age-dependent discounted life expectancy). Alternatively, age-dependent (or age- and gender-dependent) VSLY can be estimated as the ratio of the corresponding VSL and expected present value of future longevity (Aldy and Viscusi 2007, 2008, Viscusi and Hersch 2008).

Some stated-preference studies have estimated VSLY by dividing estimates of individuals' VSL by their life expectancies (e.g., Alberini et al. 2006). Only a few have estimated VSLY directly, by eliciting WTP for a specified increase in life expectancy. 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 described as being administered 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 age 60 or 70 years when the benefit was described as a risk reduction but that WTP was

[10] Hirth et al. (2000) weight future life years by expected quality of health to derive an estimate of the value per quality-adjusted life year.

about 1.6 times larger for the earlier treatment 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; respondents may not have sufficiently recognized the benefit of the earlier intervention, which is the incremental risk reduction for ages 60 to 69.

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.[11] 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 reductions 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,

[11] As noted by an anonymous reviewer, the increase in life expectancy was misleadingly characterized as the increase in area under this curve and also described as the increase in healthy life expectancy.

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 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 baseline hazard functions and their perturbations are described in the following subsection. Subsections 4.2 and 4.3 provide details of the survey instrument and describe how the survey was administered.

4.1 Baseline and perturbed hazard functions

The building blocks of the survey are baseline mortality hazard functions (for US residents of each gender and 10-year age group)[12] 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).[13]

The perturbed hazard functions are:

Transient:	$h_t(t) = h_0(t) - c$	$m \leq t < m + 10$	(7a)
	$= h_0(t)$	$t \geq m + 10$	
Additive:	$h_a(t) = h_0(t) - a$	$t \geq \text{m}$	(7b)
Proportional:	$h_p(t) = h_0(t)\ (1 - p)$	$t \geq \text{m}$	(7c)

[12] The baselines are the all race and origin hazard functions from Arias et al. (2017).
[13] For ages less than *m* the perturbed and baseline hazards are equal.

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

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.[14] 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 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 $\mu g/m^3$ or less.[15] Wright and Weinstein (1998) suggest that most cancer screening, vaccination, and other health interventions applied to the general public increase life expectancy by less than a few months.

For each increase in life-expectancy, the probability distributions of age at death for the alternative perturbations are related as mean-preserving spreads. Compared with the baseline hazard, the transient perturbation contracts the distribution of age at death, $f(t)$, and the proportional perturbation spreads the distribution. The change in standard deviation of the age at death depends on age, gender, and life-expectancy gain. For the youngest age group, individuals aged 20-29, the standard deviation of the probability distribution of age at death is 14.96 years for men and 13.97 years for women; for the oldest group, individuals aged 60-69, the corresponding values are 9.40

[14] The figures that are used to calculate VSLY are 6.9, 17.7, and 28.0 days.
[15] Relative mortality risk is estimated as 1.064 per 10 $\mu g/m^3$ (Héroux et al. 2015). US average concentration decreased from about 13 to 8 $\mu g/m^3$ between 2000 and 2020 (https://www.epa.gov/air-trends/particulate-matter-pm25-trends).

(men) and 9.75 (women). The transient perturbation with the largest life-expectancy gain (28 days) decreases the standard deviation of age at death by 0.133 and 0.142 years for men and women aged 20-29, respectively, and by 0.031 and 0.040 years for men and women aged 60-69. In contrast, the proportional perturbation with the same life expectancy gain increases the standard deviation of age at death by 0.014 and 0.008 years for men and women aged 20-29, respectively, and by 0.030 and 0.020 years for men and women aged 60-69. The additive perturbation with the same life expectancy gain decreases the standard deviation for individuals aged 20-29 (by 0.072 for men and 0.085 for women) and for women aged 60-69 (by 0.007) but increases the standard deviation for men aged 60-69 (by 0.003). Comparing the changes in standard deviation of age at death across perturbations, life-expectancy gains, genders, and the youngest and oldest age groups reveals that the difference is most strongly associated with perturbation and secondarily with age. Of the total squared variation, 23 percent is explained by perturbation type, 10 percent by age, 4 percent by life-expectancy gain, and 0.2 percent by gender.

4.2. 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 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 for all three choices and randomly selected from the three possible levels. There are 27 possible response patterns (including indifference), of which 13 are transitive.[16] 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.[17] 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.

The fourth section elicited WTP for each of three programs differing in both perturbation type (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 the perturbation type and life-expectancy gain and the order of presentation were randomized across respondents). In each case, the perturbation and increase in life expectancy were illustrated with a graphic like Figure 2 that displayed the baseline hazard, perturbed hazard, change in annual mortality risk, baseline and increase in life expectancy. 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

[16] The 27 possible response patterns can be constructed by combining the three possible preference rankings for each of the three pairs of permutations (i.e., the Cartesian product $\{T \prec A, T \sim A, T \succ A\} \times \{T \prec P, T \sim P, T \succ P\} \times \{A \prec P, A \sim P, A \succ P\}$).

[17] Responses to the three pairwise choices provide information about whether a respondent is risk averse or risk seeking over the relevant range of ages, but not about their global risk posture.

program if it were free can be interpreted as violating the assumption that the value of a life-expectancy gain is positive 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 asking 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."

4.3. 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.[18] 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

[18] 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.

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).[19]

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 all 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). 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 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 (to decrease the chance of an uninformed respondent appearing to be transitive). 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 $A \succ P \succ T$), 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

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

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.[20]

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 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

To examine how WTP depends on life-expectancy gain and the time path of risk reduction, and whether it is associated with respondents' preferences over alternative perturbations, we estimate regression models to describe WTP in the full sample and independently in each large subgroup defined by responses to the pairwise choices. In the subgroups, we estimate:

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

where w_{ijk} is individual i's annual 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-

[20] 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 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,[21] 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 specified a different perturbation type, life-expectancy gain, and initial bid. Hence estimated effects of perturbation and life-expectancy gain are identified by differences in WTP between, rather than within, respondents. These design choices help to 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 positive and suggest the respondent rejected some aspect of the valuation question (or was inattentive).

We begin with a simple test for the validity of the WTP responses: is WTP significantly larger for a greater increase in life expectancy, and is the relationship close to proportional? The first column of Table 2 reports a regression including only an intercept and the indicator variables for life-expectancy gain. The coefficients of the 7 and 28-day gains are statistically significantly negative and positive, respectively. If WTP is proportional to the life-expectancy gain, the expected values of these coefficients 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. These estimates imply that WTP is sensitive to life-expectancy gain but that the relationship is less than proportional (for life-expectancy gains between 7 and 18 days).

The second column of Table 2 reports estimates for the basic specification (equation (8)) excluding individual and survey characteristics X_i. The estimated coefficients on life-expectancy gain are nearly identical to those in the first column. Again, both are significantly different from zero and we can

[21] 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). For respondents asked about WTP per month, bids were multiplied by 12 to convert WTP to annual values.

reject the hypothesis that WTP is proportional to life-expectancy gain for the difference between 18 and 7 days, but not for the difference between 18 and 28 days.

Table 3 reports results of the basic model estimated within subgroups defined by pairwise choices over the perturbations. The estimated coefficients on life-expectancy gain differ between subsamples. 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 limited and the relevant coefficients are imprecisely estimated. Within each subgroup, consistency requires the estimated sign of the difference in WTP between each pair of perturbation types to correspond to the pairwise choice for that pair. For three subgroups (risk-seeking, risk-averse, and A≻P≻T), the signs of three differences in WTP are specified (e.g., for the risk-averse subgroup WTP for the transient perturbation should be larger than for both the additive and proportional perturbations and WTP for the additive perturbation should be larger than for the proportional perturbation).[22] Of these nine cases, the estimated difference in WTP has the correct sign in eight (the exception is that estimated WTP for the transient perturbation is smaller than for the additive perturbation for the risk-averse subgroup). The probability of having at least eight correct signs under the null hypothesis of no difference in WTP between perturbations is 0.033. However, the estimated difference in WTP is statistically significantly different from zero in only two of the nine cases (estimated WTP for the transient perturbation is statistically significantly smaller than for the additive perturbation for the risk-seeking and A≻P≻T subgroups). Consistency also requires that there should be no difference in WTP by perturbation type for the risk-neutral subgroup; the results show no statistically significant differences for this group but the standard errors are large. The full-sample results (Table 2, second column) exhibit the identical pattern of signs and significance levels as the subgroup-specific regressions (Table 3).

The coefficients of the interactions between perturbation type and the risk-neutral, risk-averse, and A≻P≻T subgroups estimate the proportional difference in WTP for different perturbations for each

[22] The difference in WTP between the additive perturbation and each of the transient and proportional perturbations is given by the estimated coefficients of the indicator variables for the transient and proportional perturbations, respectively. The difference in WTP between the transient and proportional perturbations is given by the difference between the estimated coefficients of the corresponding indicator variables.

subgroup. The six coefficients range in absolute value from near zero to one, implying that WTP and VSLY corresponding to different perturbation types differ by a factor as large as 2.7. Specifically, the ratio of the VSLY for the additive perturbation to the VSLY for the transient perturbation is about 2.7, 2.4, and 1.2 for the A≻P≻T, risk-seeking, and risk-averse subgroups, respectively.[23] Similarly, the ratio of the VSLY for the additive perturbation to the VSLY for the proportional perturbation is about 1.9 and 1.6 for the A≻P≻T and risk-averse subgroups, respectively.[24] These differences are of the same magnitude as estimates of the strength of preference for one perturbation type over another within subgroups obtained by Nielsen et al. (2010) and by Hammitt and Tunçel (2015).[25]

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 3). This suggests greater heterogeneity or more random error in these subgroups, consistent with the possibility that respondents who are less sure of their preferences or less attentive to the survey are more likely to be classified into these subgroups.

Tables 4 and 5 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 2 and 3). Estimates for the full sample (Table 4) 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 5) 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

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[23] WTP for the additive perturbation is estimated as exp(1.005) ≈ 2.7, exp(0.876) ≈ 2.4, and exp(0.213) ≈ 1.2 times as large as for the transient perturbation.

[24] WTP for the additive perturbation is estimated as exp(0.654) ≈ 1.9 and exp(0.457) ≈ 1.6 times as large as for the proportional perturbation.

[25] Nielsen et al. (2010) and Hammitt and Tunçel (2015) elicited respondents' strength of preference between alternative perturbation types using follow-up questions to the pairwise choices between perturbations having equal life-expectancy gains. In the follow-up questions, respondents chose between the initially dispreferred perturbation and the initially preferred perturbation type with a smaller life-expectancy gain. Nielsen et al. (Table 4) report that respondents were on average indifferent when the life-expectancy gain of the initially preferred perturbation was decreased by 1.5 to 2.3 months (of an initial 6 month gain), which implies indifference to perturbation type with a ratio of life-expectancy gains of 1.3 to 1.6. From Hammitt and Tunçel (Table 2) we estimate the median respondent was indifferent to perturbation type with a ratio of life-expectancy gains of 2.4.

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. This is further evidence that the results can be interpreted as valid measures of WTP for reductions in mortality risk. 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 although estimated WTP in the risk-neutral subgroup is significantly smaller in 2021 than in 2020 (Table 5). The effect of taking an oath to respond truthfully is small and not significantly different from zero in the full sample (Table 4); the estimated coefficients by subgroup are inconsistent in sign and magnitude, including significantly negative in the risk-neutral and significantly positive in the intransitive subgroup (Table 5). There is no evidence that oath taking improves validity in the sense of enhancing sensitivity to scope: adding the indicator variable for oath taking and its interactions with the indicators for life expectancy gain to the simple model reported in the first column of Table 2 reveals no significant effects (results not shown).

In summary, estimated WTP is significantly associated with the magnitude of the life-expectancy gain, though the estimated effect is less than proportionate. On average, there is no significant difference in WTP by perturbation type, although there is evidence of differences that are consistent with respondents' preference ordering over the perturbation types: the risk-seeking and A≻P≻T subgroups express significantly smaller WTP for the transient (least-preferred) risk reduction and eight of the nine differences in WTP between perturbation types by subgroup are of the correct sign ($p < 0.05$). The variation of WTP with life-expectancy gain and the lack of variation in annual WTP between questions that elicited annual or monthly WTP support the validity of these results as estimating WTP for mortality-risk reduction.

5.3. Estimated VSLY

Estimates of WTP from the regressions specified by equation (8) are sensitive to the method used for retransformation of 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 (second column of Table 2). Using this value, $exp(\hat{\sigma}^2/2) \approx 200$.

Given the large difference between predicted mean and median WTP, we supplement our primary model with a linear regression identical to equation (8) (excluding X_i) but using WTP (w_{ijk}) as the dependent variable. Although the linear model implies (counterintuitively) that differences in WTP between perturbation types are independent of the life-expectancy gain, it is not subject to the retransformation problem. Results for the linear model are presented in the appendix (Table A). 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 A≻P≻T subgroup with the proportional perturbation, these effects are also statistically significant in the parallel logarithmic model (the second column of Table 2).

To estimate VSLY, we predict mean annual WTP for each perturbation and life-expectancy gain using the full-sample regression model (Table 2, second column), multiply by 10 to produce total WTP (because respondents are told they would pay for 10 years[26]) and divide by the life-expectancy gain (measured in years).

Table 6 presents the calculated VSLY. The first three rows show the results for each life-expectancy gain and subgroup (averaging across perturbation type). The following row shows the geometric mean of the values by life-expectancy gain and the final row shows the corresponding results from the linear model. 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 3). Estimates from the risk-neutral subgroup are also less compelling than for the other transitive subgroups given that some of the respondents in this group may have indicated indifference between perturbations as an expression of uncertainty; the residual standard deviation for the risk-neutral subgroup is larger than for each of the other transitive subgroups (Table 3).

Among the transitive subgroups, VSLY ranges from about \$300,000 to \$900,000 across life-expectancy gains (from \$1.1 to \$2.4 million for the intransitive subgroup). Because predicted WTP is less than proportional to life-expectancy gain, the calculated VSLY is smaller for larger life-expectancy gains. Averaging across transitive subgroups, the values calculated for the 18 and 28-day

[26] 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.

gains in life expectancy are close ($429,000 and $400,000) because WTP is close to proportional for these gains, but VSLY is much larger ($833,000) for the 7-day life expectancy gain.

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 transitive subgroups, this summary ranges from $387,000 to $570,000 and averages $523,000. The variation by subgroup is modest; VSLY is always within 10 percent of the transitive subgroup average, except the value for the risk-neutral subgroup is about 25 percent smaller. In contrast, the value for the intransitive subgroup is almost three times larger. For comparison, calculated VSLY using the linear model (the last row of Table 6) is modestly smaller than using the logarithmic model, about 25 percent averaging across transitive subgroups.

These estimates of VSLY are modestly 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. Using hedonic-wage methods (and discounting future life years at 3 percent annually) Aldy and Viscusi (2008) estimated an average worker VSLY of about $520,000 and Viscusi and Hersch (2008) estimated an average VSLY for male and female smokers of about $680,000 and $490,000, respectively.[27] If future life years were not discounted, these values would be roughly 40 percent smaller (assuming life expectancy near 40 years), i.e., about $300,000 for all workers and $400,000 and $300,000 for male and female smokers, respectively.

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) either for a population or an individual. 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.

[27] Aldy and Viscusi (2008) report VSLY of about $300,000 and Viscusi and Hirsch (2008) report VSLY of $390,000 and $280,000 for male and female smokers, respectively, in 2000 dollars. To convert to 2021 dollars we adjust for inflation using the CPI for all urban consumers (multiply by 1.54) and for mean income growth from the Current Population Survey assuming an income elasticity of 1.0 (multiply by 1.13).

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 similar to those found in previous studies (Nielsen et al. 2010, Hammitt and Tunçel 2015). 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 of transitive patterns 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, estimated WTP depends on life-expectancy gain 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, eight of the nine differences in WTP between perturbation types for the subgroups for which differences are predicted are of the correct sign, a result that allows us to reject the hypothesis of no concordance between the difference in WTP and the pairwise choice between perturbation types. WTP for the transient (least-preferred) perturbation is significantly smaller than for the other perturbations in the risk-seeking and A≻P≻T subgroups.

Our estimates of VSLY are comparable to, but somewhat larger than, conventional estimates obtained by dividing estimated VSL for a current risk by average life expectancy. Estimates based on predicted mean WTP range between $387,000 and $570,000 in the transitive subgroups, with an overall mean of $523,000. If respondents discount future payments at an annual rate of 3 or 5 percent, these values would be 5 or 13 percent smaller, respectively. 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.

Because most environmental, health, and safety regulations that decrease mortality risk do so for an extended period, estimates of the value of a continuing risk reduction are more relevant than conventional estimates of the value of reducing risk for a single year. Moreover, because the value of any change to an individual's hazard function can be characterized as a VSL or VSLY, estimates of the two 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 these 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 2. Estimated log(WTP), basic model

	LE only	Basic model
Transient		0.007
		(0.257)
Proportional		0.040
		(0.255)
LE 7	-0.281*	-0.279*
	(0.183)	(0.183)
LE 28	0.380**	0.388**
	(0.179)	(0.178)
RS		0.599
		(0.389)
RA		0.503
		(0.447)
A≻P≻T		0.666
		(0.434)
Other transitive		0.373
		(0.272)
Intransitive		1.361***
		(0.318)
Trans*RS		-0.876*
		(0.465)
Prop*RS		0.005
		(0.456)
Trans*RA		-0.213
		(0.563)
Prop*RA		-0.457
		(0.559)
Trans* A≻P≻T		-1.005*
		(0.543)
Prop* A≻P≻T		-0.654
		(0.532)
Intercept	4.756***	4.347***
	(0.128)	(0.299)
σ	3.278***	3.250***
	(0.065)	(0.065)
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 3. Estimated log(WTP) by subgroup, basic model

	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 4. 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 5. 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 6. Calculated VSLY by life-expectancy gain ($000)

	Risk neutral	Risk seeking	Risk averse	A≻P≻T	Other transitive	Intransitive	Mean	Mean (transitive)
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
Geometric mean	387	570	520	473	562	1,511	636	523
Linear model	430	364	384	369	419	550	411	393

Notes: Values by LE are arithmetic means over perturbation type. Last two columns are means of corresponding rows (weighted by subgroup frequency). Last two rows are arithmetic and geometric means of the three values by LE.

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

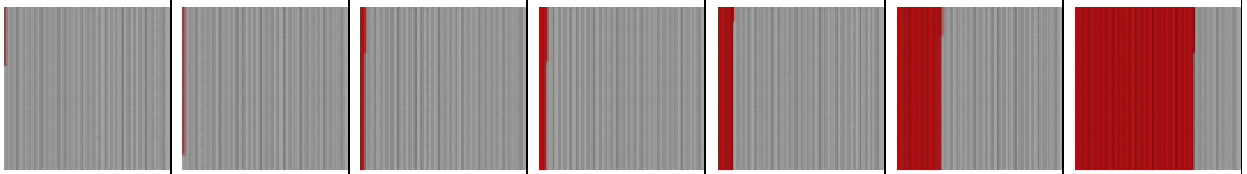


Figure 1. Introduction to hazard function and life expectancy (woman aged 20-29 years).

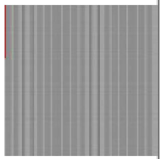
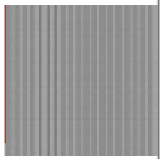
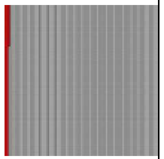
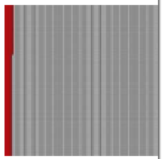
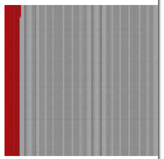
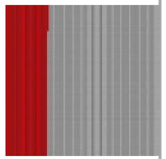
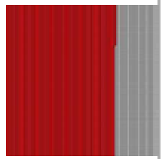
	Future Decades	30 – 39 Years	40 – 49 Years	50 – 59 Years	60 – 69 Years	70 – 79 Years	80 – 89 Years	90 – 99 Years	Life expectancy
	Mortality risk without the reduction (out of 100,000)	585 	1 145 	2 643 	5 615 	12 153 	30 239 	68 075 	**54 years**
Program X	**Mortality risk reduction (out of 100,000)**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

Appendix. Linear WTP model

Table A. Estimated WTP

Transient	17.4
	(196.8)
Proportional	213.9
	(195.6)
LE 7	-55.8
	(138.5)
LE 28	186.8
	(137.5)
RS	-39.7
	(294.4)
RA	-94.3
	(340.1)
A>P>T	297.3
	(329.1)
Other	-43.8
	(205.0)
Intransitive	494.3**
	(246.1)
Trans*RS	-481.4
	(353.8)
Prop*RS	-218.9
	(351.8)
Trans*RA	-52.8
	(427.5)
Prop*RA	-226.5
	(427.4)
Trans*A>P>T	-806.7**
	(408.3)
Prop*A>P>T	-838.2**
	(406.7)
Intercept	1655.7***
	(224.7)
σ	2770.5***
	(44.5)
N	2676

Notes: *, **, *** denote significantly different from zero at 10, 5, 1 percent, respectively.