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Does the value of a statistical life vary with age and health status? Evidence from the US and Canada☆

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Abstract

This paper provides an empirical assessment of the effects of age and baseline health on willingness to pay (WTP) for mortality risk reductions by reporting the results of two contingent valuation surveys: one administered in Hamilton, Ontario and the other to a national sample of US residents. Respondents for both surveys were limited to persons aged 40 years and older to examine the impact of age on WTP. Using the WTP responses and those regarding respondent's own and family health histories, we find weak support for the notion that WTP declines with age, and then, only for the oldest respondents (aged 70 or above). Furthermore, we find no support for the idea that people with chronic heart or lung conditions, or cancer,

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are willing to pay less to reduce their risk of dying than people without these illnesses. If anything, people with these illnesses are willing to pay more.
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1. Introduction

In the public health literature, the benefits of life-saving programs are measured in terms of quality-adjusted life years (QALYs), a measure that explicitly weights reductions in premature mortality by remaining life expectancy and by the health status of the persons saved. In benefit–cost analyses of environmental programs, where reductions in premature mortality are valued by what people would pay for them, there is considerable controversy as to whether willingness to pay (WTP) for mortality risk reductions should vary with the age and health status of the persons whose lives are extended.[1]

The answer to these questions has important implications for policy: according to epidemiological studies, the majority of statistical lives saved by environmental programs appear to be the lives of older people and people with chronically impaired health [14–16]. There are two reasons why older persons are likely to benefit disproportionately from reductions in pollution. First, epidemiological studies typically assume that the effects of a change in exposure are proportional to baseline mortality [13,14]. Since persons over 65 account for three-quarters of all deaths in the US and Canada, a larger proportion of statistical lives will be saved among the old than among the young. Second, some epidemiological studies have found larger changes in mortality rates for people over 64 than for younger people [15,16]. Epidemiological studies also suggest that persons with chronic heart or lung conditions are likely to benefit disproportionately from improvements in air quality [14,15,17].

It has been conjectured that older people should be willing to pay less for a fatal risk reduction than younger people on the grounds that they have fewer expected life years remaining. Indeed, some economists have argued that the value of a statistical life (VSL) should be converted to a value per statistical life year (VSLY), and that lives saved should be valued by multiplying remaining life expectancy by the VSLY [12]. This procedure implies that each year of life is equally valuable, and that the VSL is proportional to remaining life expectancy. Whether this approach is consistent with welfare economics, however, depends on how, empirically, WTP for a reduction in risk of death varies with age.

It has also been argued that people in ill health should be willing to pay less for a fatal risk reduction because their utility from an additional year of life is less than that of healthy people. In the QALY literature, saving the life of a chronically ill person is valued less than that of a person in good health [7]. This argument has been used to assign lower VSLs to beneficiaries of air pollution control programs than that currently used by USEPA [5]. There is no published

[1] This controversy recently spilled over into the public arena, when elderly people demonstrating at EPA "listening sessions" dubbed EPA's practice of using VSLs for elderly populations in alternative analyses that are lower than those typically used in primary analyses the "senior discount." Administrator Whitman renounced the procedure shortly thereafter [19].

empirical evidence, however, showing that people in poor health would pay less than healthier individuals to reduce their risk of dying.

This paper provides an empirical assessment of the effects of age and baseline health on WTP for mortality risk reductions by reporting the results of two contingent valuation surveys designed to test the above hypotheses. One survey was administered to residents of Hamilton, Ontario and the other to a nationally representative sample of US residents. Both surveys elicited respondents' WTP for reductions in mortality risk of different magnitudes. Respondents were limited to persons aged 40 years and older, including those older than 70, to examine the impact of age on WTP. Extensive information was collected about each respondent's health status to see if it systematically influences WTP.

Our results provide weak support for the notion that WTP declines with age, but only after age 70. Specifically, in our Canadian sample, WTP declines by about 30% after age 70 compared to WTP at younger ages. The US sample shows no such statistically significant decline. We also find no support for the idea that people with chronic heart or lung disease, or cancer, are willing to pay less to reduce mortality risk than people without these illnesses. If anything, people so stricken are willing to pay more.

2. How should age and health status affect WTP for mortality risks?

Several authors have used the life cycle consumption model with uncertain lifetime to [22] derive an expression for an individual's WTP for a reduction in his risk of death [4,18]. In this model, a person at the beginning of period j (i.e., at age j) receives expected utility of V_j over the remainder of his lifetime:

$$V_j = \sum_{t=j}^{T} q_{j,t}(1+\rho)^{j-t} u_t(C_t), \tag{1}$$

where V_j is the present value of utility of consumption in each period, $u_t(C_t)$, times the probability that the individual survives to that period, $q_{j,t}$, discounted to the present at the subjective rate of time preference, ρ. T is the maximum length of life. V_j is maximized subject to initial wealth, W_j, and a budget constraint that reflects opportunities for borrowing and lending. The two cases usually considered are the case of actuarially fair annuities and the more realistic situation in which the individual can borrow and lend at the riskless rate r, but can never be a net borrower, implying that he must have non-negative wealth in all periods

$$W_t = W_j + \sum_{k=j}^{t}(1+r)^{j-k}(y_k - C_k) \geqslant 0, \quad j \geqslant t \geqslant T, \tag{2}$$

where y_k is income at age k.[2]

The life-cycle model can be used to determine the amount of initial wealth that an individual would give up to reduce D_j, the probability that he dies during the current period. A reduction in

[2] We focus on the more realistic case in which the individual can never be a net borrower. Allowing the individual to purchase actuarially fair annuities would not, however, change the basic conclusions of this section—that the impact of age and health status on WTP for a reduction in risk of death is, in general, ambiguous.

D_j will increase the probability that the person survives to all future periods since, by definition, $q_{j,t}$ is the product of the probabilities that the individual does not die in all periods from j to $t-1$

$$q_{j,t} = (1 - D_j)(1 - D_{j+1})\ldots(1 - D_{t-1}). \tag{3}$$

The rate of substitution between D_j and W_j corresponds to the VSL for a person of age j, VSL_j,

$$VSL_j = (\partial V_j/\partial D_j)/(\partial V_j/\partial W_j) = dW_j/dD_j. \tag{4}$$

The amount an individual is willing to pay for the change in D_j is, in turn, the product of the VSL_j and the size of the risk reduction[3]

$$WTP_j = (VSL_j)dD_j. \tag{5}$$

If the individual's wealth constraint (2) is binding only at T, an expression for VSL_j may be derived by appending the constraint $W_T = 0$ to (1) and using the Envelope Theorem to evaluate (4).[4] The resulting expression, (6), implies that VSL_j may be written as the product of the reciprocal of the probability that the individual survives the current period, $(1 - D_j)^{-1}$, times the present value of expected utility of consumption from period $j+1$ onward, converted to dollars by dividing by the marginal utility of consumption, $\partial u_j/\partial C_j$.

$$VSL_j = (1 - D_j)^{-1} \sum_{t=j+1}^{T} q_{j,t}(1+\rho)^{j-t} \frac{u_t(C_t)}{\partial u_j/\partial C_j}. \tag{6}$$

How should the VSL change with age j? The first term in (6), $(1 - D_j)^{-1}$, unambiguously increases with age: As people age, their probability of surviving the current period falls and, for this reason, their WTP to reduce their risk of death should increase.[5] How the present value of expected utility of consumption—the remainder of the equation—changes with age is ambiguous. If utility of consumption $u_t(C_t)$ were constant over time, then the present value of expected utility of consumption would be proportional to discounted remaining life expectancy, $\sum_{t=j+1}^{T} q_{j,t}(1+\rho)^{j-t}$. The latter unambiguously decreases with age (j), and motivates the hypothesis that WTP for mortality risk reductions should fall with age. In general, however, $u_t(C_t)$ is not constant. In fact, it may plausibly increase with t either because C_t is increasing as one ages, or because the enjoyment derived from C_t is higher as one ages.[6]

If health status is treated as exogenous, it can easily be incorporated into Eq. (6).[7] Let H_t represent health status at age t, with lower values indicating poorer health status. It is certainly

[3] This is an approximation to the individual's WTP, defined as the amount one can subtract from W_j and keep expected utility constant when D_j is altered.

[4] Formally, $L_j = \sum_{t=j}^{T} q_{j,t}(1+\rho)^{j-t}u_t(C_t) + \lambda[W_j + \sum_{t=j}^{T}(y_t - C_t)(1+r)^{j-t}]$. By the Envelope Theorem, $(\partial V_j/\partial D_j)/(\partial V_j/\partial W_j) = (\partial L_j/\partial D_j)/(\partial L_j/\partial W_j)$. Using the fact that $\lambda = \partial u_j/\partial C_j$ from the first-order conditions yields Eq. (6).

[5] This is certainly true in the US and Canada for persons over age 40.

[6] Fernandez-Villaverde and Krueger [6] find that consumption expenditure per adult equivalent household member peaks at about age 50 in the US, based on data from the 1980–1998 Consumer Expenditure Surveys. Indeed, consumption at age 50 is about 50% higher than at age 70. Consumption may in fact be more enjoyable when one is no longer raising children or caring for one's parents. The effect of age on WTP is ambiguous.

[7] If health status were endogenous, one would have to examine the impact of a change in D_j on health expenditures.

reasonable that H_t is inversely related to the conditional probability of dying at age t, D_t, and that it affects the utility of consumption at age t. Will a person with poorer current health (lower H_j) be willing to pay more or less than a person in better health for a reduction in D_j? If lower values of H_j signify poorer health, then the first term on the right-hand side of (6) will be higher the lower is H_j. Persons in poorer health presumably have smaller chances of surviving the current period and, for that reason, should be willing to pay more to reduce D_j. The effect of health status on the rest of the equation is ambiguous. Even if discounted remaining life expectancy is lower for those with lower current health status, one can say little about their time pattern of consumption, or about the way in which health affects the marginal utility of consumption. The implications of this section are that little can, in general, be said about the impact of current age (j) or current health status (H_j) on the value of mortality risk reductions.

3. The surveys

3.1. The commodity valued

The goal of our surveys is to estimate individuals' WTP for a reduction in their conditional probability of dying during the current period (D_j). Periods are treated as 10 years long.[8] After being told the baseline risk of death over the next 10 years for someone of their race and gender, individuals are asked whether they would purchase a product (not covered by health insurance) that would reduce this risk by either 1 in 1000 (1 in 10,000 annually) or 5 in 1000 (5 in 10,000 annually), at a stated price.[9] Payment for the commodity is to be made annually, over 10 years.

Both the baseline risk of death and the risk reductions are communicated graphically. Baseline risk of death over the next 10 years is represented by red squares on a white grid containing 1000 squares. Reductions in risk of death are shown by turning the appropriate number of red squares to blue.

3.2. The structure of the survey

Our survey begins with questions about the respondent's health history and the health history of his family. This is followed by exercises that acquaint the respondent with the concept of risk and test his comprehension. Subjects are introduced to simple probability concepts using coin tosses and roulette wheels, working up to our standard risk communication device—a 1000-square grid in which risks are represented using red squares. To test their comprehension, respondents are asked to compare grids for two hypothetical people (person A and person B) and to determine which of the two has the higher risk of death. They are also asked to select which of the two people they would rather be. The baseline risk of death for a person of the respondent's age and gender is then presented both numerically and graphically.

[8] In focus groups, individuals more readily accepted information about their risk of dying and changes in their risk of dying over a 10-year period than over shorter periods.

[9] We chose a risk reduction of this order of magnitude because risk changes valued in labor market studies are typically of this size and risk reductions of comparable size are often delivered by environmental programs.

Table 1
Bid structure in the Canada and US mortality risk survey (1999 Canadian $ and 2000 US $ respectively)[a]

Group of respondents	Initial payment question	Follow-up question (if "Yes")	Follow-up question (if "No")
I	100 (70)	225 (150)	50 (30)
II	225 (150)	750 (500)	100 (70)
III	750 (500)	1100 (725)	225 (150)
IV	1100 (725)	1500 (1000)	750 (500)

[a] Bids in the US version of the survey in parentheses. The bids used in the US study were obtained from the bids used in the Canada study after conversion to US dollars using purchasing power parity.

It is sometimes argued that respondents in contingent valuation surveys find it difficult to report their WTP for a mortality risk reduction because they are not accustomed to trading income for reduced risks. To mitigate this problem, we first acquaint respondents with *quantitative* risk reductions resulting from medical tests and products that are likely to be familiar to the respondent (e.g., mammograms, colon cancer screening tests, medicine to reduce blood pressure). In doing so, we provide only qualitative cost information for each action or product ("inexpensive," "moderate" and "expensive").

This is followed by the WTP questions. Information about WTP is obtained through a combination of dichotomous choice payment questions with follow-ups, and open-ended questions. Respondents are asked an initial dichotomous choice question: would you buy the product at a price randomly chosen from one of four pre-determined values? (See Table 1 for bid values.) Those who answer yes are asked if they would pay a higher price. Those who answer no are asked if they would pay a lower price. Respondents giving "yes–yes" or "no–no" responses are asked a final open-ended question.

Each respondent was asked his WTP for two risk reductions. Respondents in Wave 1 were asked to value the 5-in-1000 risk reduction first, whereas those in Wave 2 were asked to value the 1-in-1000 risk reduction first. After each question, respondents were asked to indicate their degree of certainty about the WTP responses on a scale of 1–7. Since the respondent's understanding of risks and interpretation of the scenario can affect WTP, we included debriefing questions at the end of the questionnaire to identify respondents who had trouble comprehending the survey or who did not accept the risk reduction being valued. These were followed by questions about the respondent's income, and Short Form-36 (SF-36), a questionnaire used in medical research to assess mental and physical health [21].[10]

[10] The survey instruments used in Canada and the US were very similar. In particular, the WTP questions remained unchanged from one venue to the next. Minor wording changes were made in other parts of the questionnaire based on our experience in Hamilton. Also, several questions were reworded in the US survey to elicit more detail on health history and demographics. The Canadian survey included voice-overs for all computerized questions whereas their use

3.3. Administration of the questionnaire

The Canadian survey was self-administered using a computer by 930 residents of Hamilton, Ontario. Subjects were recruited by telephone through random-digit dialing and asked to take the survey at a facility in downtown Hamilton.[11] The survey took place over 5 months in the spring of 1999.

The respondents in the US survey were reached through a technology called Web-TV that involves attaching a special device (resembling a cable box) to a television. A remote control device or a keyboard enables the user to access the Internet, using the television as a monitor. Knowledge Networks recruits individuals to participate as panel members in exchange for the technology and free Internet access. The panel members are recruited by telephone using random-digit dialing, and are representative of the US population in terms of gender, age, race and income. Panel members are randomly selected to complete surveys.[12] Knowledge Networks administered our survey to a randomly selected sample of their panel members fitting our age profile in August 2000. One thousand eight hundred persons were contacted to take the survey, and within three days 1200 of them (our target sample size) had completed the survey.

4. Sample characteristics and responses

4.1. Characteristics of the respondents

The respondents in our Canadian and US surveys are somewhat different since they were sampled from different populations. In this section we describe our respondents, focusing on their age and health. Table 2 describes features of each sample other than health. It shows that both samples were well balanced in terms of gender, with each having slightly more women than men. While respondents in the Canadian study were all Caucasian, the US sample did include blacks (11%) and Hispanics (8%).

Due in part to differing racial compositions, the baseline risks reported to respondents were different in the two studies.[13] The average baseline risk was 123 in 1000 in the Canadian study, and 187 in 1000 in the US study. Blacks, included in the US sample, tend to have higher baseline risks—except when very old—compared to whites. The US sample also included respondents over 75 whereas the Canadian sample did not. When these elderly respondents and blacks are excluded from the US sample, baseline risks decline substantially and are comparable to those for the Canadian sample.

(*footnote continued*)
in the US version was seriously curtailed due to technological constraints.

[11] Because of the need to travel to a centralized facility, response rates were low. Out of 17,841 residential phone contacts 8260 were "cooperative," but 4917 households proved ineligible for age reasons. Among the 3591 eligible households, 455 declined to participate because of mobility problems and 1079 refused, stating that the incentive payment (C$35) was insufficient. 1545 persons agreed to participate in the survey, but only 930 (60%) kept their appointments. The response rate, calculated as the number of respondents successfully completing the study (930) divided by the number of eligible contacts (3591), is therefore 26%.

[12] More information about Knowledge Networks is available on their website: www.knowledgenetworks.com.

[13] The baseline risks presented to respondents in each study reflected country specific estimates by age–race–gender.

Table 2
Comparison across Canada and US mortality risk studies: characteristics of respondents

Characteristic	Sample average or percent of sample	
	Canada	US
Age	54.2 years	54.4 years
Male	46%	47%
Racial and ethnic composition		
African-American	—	11%
Hispanic	—	8%
White	100%	82%
Baseline risk of dying over the next 10 years		
Entire sample	123	187
African-American	—	174
No African-Amer. or persons older than 75	—	147
Household characteristics		
Annual household income (US $)		
Mean	$46,800	$53,000
Median	$50,000	$55,000
Years of schooling	13.7	13.0
Married	—	72%
Household size	—	2.6
Number of adults in the household	—	2.2
Percent urban/suburban in county of residence	100%	78%

Average household incomes are similar in both studies, as are years of schooling. The US study included participants from areas that are classified as neither urban nor suburban (22% of the sample), while the Canadian study, by design, covers only residents of the urban and suburban area of Hamilton.

4.2. Respondent health

We obtained a variety of information about respondent health. First, we asked respondents to rate their health compared to others the same age. We then asked if they had been diagnosed with various chronic illnesses (e.g., asthma, chronic bronchitis, emphysema, heart disease, cancer, high blood pressure, and stroke). We also inquired about time spent in hospital. Finally, we administered all SF-36 questions.

Table 3 displays descriptive statistics for respondent health status. Because participants in Canada had to be well enough to travel to a centralized facility to take the survey, this sample is likely to be relatively healthy. By contrast, US respondents participated from their homes, allowing the inclusion of less healthy individuals and persons with impaired mobility. The difference in health status across the samples is borne out in the table. A higher percentage of

Table 3
Comparison across Canada and US mortality risk studies: Health status

Health condition	Sample mean or percent	
	Canada	US
Respiratory Illness		
Has asthma	—	10%
Has bronchitis	—	7%[a]
Has emphysema	—	4%
Has one or more of these conditions	14%	16%
Heart disease		
Has angina pectoris	—	8%
Has had a myocardial infarction (heart attack)	—	8%
Has coronary disease	—	7%
Has one or more heart condition	10%[b]	21%
Other		
Has had a stroke	—	4%
Has been diagnosed with cancer	3%	11%
ER visit in last 5 years or hospitalization in last year for ongoing heart or lung problems	12%	12%
Has family history of cancer	49%	52%
Has family history of chronic illness (excluding cancer)	79%	50%
Has no medical insurance[c]	31%	6%
Rates own health as good or excellent, relative to others of the same age	57%	53%
SF-36 Scores		
General health	70	67
Physical functioning	81	78
Vitality	63	59
Role-emotional	81	87
Mental health	76	77

[a] Chronic bronchitis.
[b] Heart disease.
[c] No medical insurance is defined in Canadian sample as "no supplemental insurance coverage."

respondents in Hamilton described themselves as having good or excellent health relative to others the same age (57.2%) compared to US respondents (53.1%). Furthermore, the fractions of the sample with various types of chronic illness were higher in the US.[14] While 3.4% of the Hamilton respondents said they had been diagnosed with cancer (a figure in keeping with local health statistics), 11% of the US sample reported to have been diagnosed with cancer.

[14] The questions on chronic illness were asked differently in the two studies. Canadian respondents were asked, in a single question, whether they were ever diagnosed with one or more of the following illnesses: asthma, bronchitis, or emphysema. Respondents in the US, however, were asked whether they had ever been diagnosed with the following illnesses in four distinct questions (i) asthma, (ii) chronic bronchitis, (iii) emphysema, or (iv) other respiratory illnesses. Similarly, the respondents in Hamilton were asked if they had been diagnosed with heart disease, whereas the subjects in the US were asked four questions—whether they had ever been diagnosed with (i) angina pectoris, (ii) coronary heart disease, (iii) other heart disease, and (iv) heart attack.

Table 4
Comparison across Canada and US mortality risk studies: probability comprehension

	Percent of the sample	
	Canada	US
Probability test questions answered incorrectly		
1st probability test question	11.6	12.2
2nd probability test question (FLAG4)	1.1	1.8
Indicates preference for individual with higher risk of death in		
1st probability choice question	13.0	10.8
Follow-up "confirmation" question (FLAG5)	1.3	1.3
Other indicators of probability comprehension		
Fails both probability test and choice questions (FLAG1)	2.6	3.7
Claims to understand probability poorly (FLAG6)	7.0	16.2

4.3. *Probability comprehension and acceptance of the scenario*

Valuing risk reductions through direct questioning techniques require that subjects understand probabilities and accept the WTP scenarios presented to them. This section summarizes respondents' comprehension of probabilities and reports the percent of respondents who questioned the survey assumptions. Respondents who failed our probability tests were dropped from the analysis. Responses to debriefing questions indicating that the respondent did not believe his baseline risk of death or had doubts about product effectiveness were used to create dummy variables to test whether people who questioned some aspect of the questionnaire had significantly different WTP from those who did not.

To test their comprehension of probabilities, we asked respondents several questions using side-by-side grids of squares to convey the chances of dying for two people, person A and person B. The first question, the *probability test* question, asks which of the two people has the higher probability of dying. The second question, the *probability choice* question, asks which of the two people the respondent would rather be. If respondents answer the probability test question incorrectly, another explanation of the concept is provided and the respondent is asked a second probability test question. Should respondents indicate a preference for the person with the higher risk in the choice question, they are asked to confirm their selection following an additional explanation of the grids.

Table 4 reports the results of the probability test and choice questions. The table shows that roughly 12% of respondents answered the initial probability test incorrectly. Following an explanation of the error, a much smaller portion persisted in providing an incorrect answer in the second test question (1.1% in Canada and 1.8% in the US). A similar proportion initially indicated a preference for the person with the higher risk in the probability choice question (11% in the US and 13% in Canada). However, most of these respondents corrected their answer when asked the question a second time. Only 1.3% of each sample confirmed their preference for being the person with the higher risk of death.

Table 5
Comparison across Canada and US mortality risk studies: acceptance of the product and scenario

	Percent of the sample	
	Canada	US
Did not believe the risk figures (FLAG7)	19.7	24.5
Thought own risks were higher	15.9[a]	20.5[a]
Thought own risks were lower	84.1[a]	79.5[a]
Doubts effectiveness of the product/action (FLAG8)	30.6	33.5
Doubts about effectiveness influenced WTP (FLAG9)	19.7	21.1
Thought about possible side effects of the product (FLAG10)	25.0	15.4
Thought of other benefits of the product (FLAG11)	48.7	36.6
Other benefits to self	—	39.7[b]
Benefits to other people of living longer	—	25.2[b]
Improved health for other people	—	25.7[b]
Did not consider whether he could afford the product/action (FLAG15)	26.0	37.4
Did not understand the timing of the payments (FLAG16)	13.0	14.0

[a] Percent of the respondents who did not believe the risk figures.
[b] Percent of the respondents who thought of other benefits of the product.

Combining the responses to the test and choice questions, about 3% of each sample answered the initial probability test question incorrectly *and* indicated a preference for the person with the higher chance of dying in the choice question. These respondents have been removed from subsequent analyses.

In Table 5 we examine the acceptance of the risk-reducing product/action and the scenario presented in the questionnaire. Roughly 20–25% of the respondents did not believe the baseline risk figures that were presented to them. Most of these respondents thought that their own risks of death were lower than the questionnaire stated. Approximately one-third of the respondents in each sample had doubts about the effectiveness of the product/action, with a large fraction of these respondents stating that these doubts influenced their WTP for the product/action itself.

Some respondents worried about side effects of the product, while others thought that the product would yield additional benefits. A larger proportion of Hamilton respondents were concerned about possible side effects and admitted thinking about other benefits of the product than participants in the US. In the US study, respondents were asked about the kinds of additional benefits they had in mind. Other benefits for these respondents included additional benefits to themselves (40%), benefits to other people (e.g., family members) of their living longer (25%) and improved health for other people (26%).

In Hamilton, 26% of respondents noted that they did not consider whether they could afford the product/action when answering the payment question. In the US, the fraction of the sample reporting such behavior was even higher (37%). As discussed in Krupnick et al. [11], these responses were common among people who were not willing to pay anything for the product. We conclude that most of these respondents had already ruled out the purchase of the product making the price irrelevant to their decision. Finally, about 14% of respondents revealed that they had not understood that they would be required to make payments every year for 10 years to receive the product and its risk-reducing benefits.

5. Willingness to pay results

5.1. Validity of WTP responses

Before we examine the impact of age and health on WTP, it is important to establish criteria that WTP responses must satisfy for consistency with economic theory. We use three such criteria. First, the percentage of respondents answering "yes" to the initial payment question must decline with the dollar amount presented to respondents. Second, respondents should be willing to pay more for a larger risk reduction. Third, under the assumptions in Section 2, WTP should be proportional to risk reduction size.[15]

To test if the first criterion is met, we use responses to the initial payment question. Figs. 1 and 2 show that the first requirement is easily met in both studies: the percentage of "yes" responses clearly declines with the bid amounts used for the initial payment questions. For the 5 in 1000 risk reduction (wave 1), about 73% of the respondents are willing to pay the lowest bid amounts used (C$100 in Canada and $70 in the US). Smaller proportions are willing to pay the highest bid amounts (26% in Canada and 35% in the US).[16] A similar reaction to the 1 in 1000 risk reduction (wave 2) is also found in both countries. The fraction of the sample willing to pay for the 1 in 1000 risk reduction declines from 49% (for C$100) to 20% (C$1100) in Canada, and from 44% (for $70) to 13% ($725) in the US.[17]

Furthermore, the percentages of respondents willing to pay for the 1 in 1000 risk reduction are smaller than those for the 5 in 1000 risk reduction at all bid levels. We expect that, when a formal estimate of mean WTP is obtained, WTP for the two reductions will be found to be significantly different.

5.2. External scope tests

To test whether WTP increases in proportion to risk reduction size we utilize the responses to the initial payment questions as well as the subsequent round of follow-ups, fit a formal statistical model of WTP, and use the latter to produce estimates of mean WTP for the specified risk reductions.

A key decision is whether or not to use the open-ended responses that follow the first and second dichotomous choice questions. In both surveys, subjects answering "no" to both

[15] Eq. (4) implies that WTP can be approximated by the rate at which an individual is willing to trade wealth for risk (VSL_j), multiplied by the size of the risk reduction. For small variations in D_j, VSL_j should remain constant.

[16] At each bid value, we performed chi square tests to check for significant differences between the proportions of "yes" responses in the Canada and US studies. We did not find any significant differences. The chi square values range between 0.02 and 2.89, and thus fall in the acceptance region of the chi square with one degree of freedom.

[17] Chi square tests indicate that there are no significant differences in the proportions of "yes" responses to the initial bid values across the two studies.

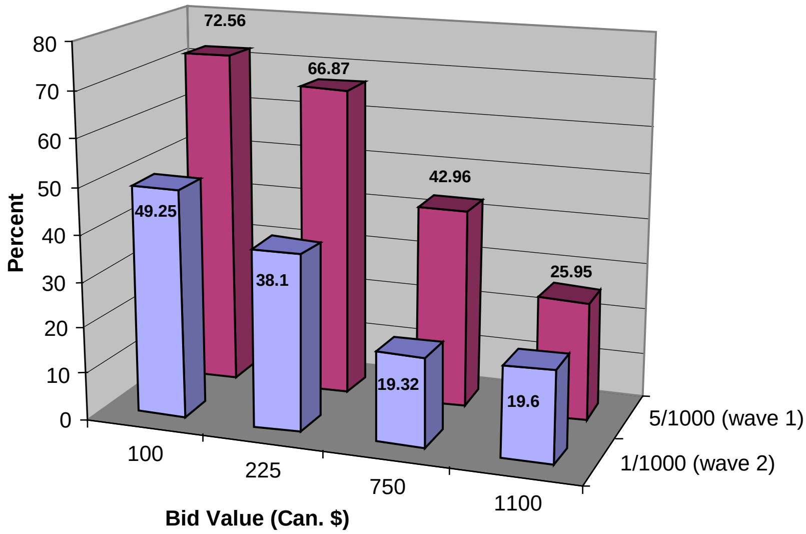

Fig. 1. Percent of "yes" responses by bid value: Canada Study.

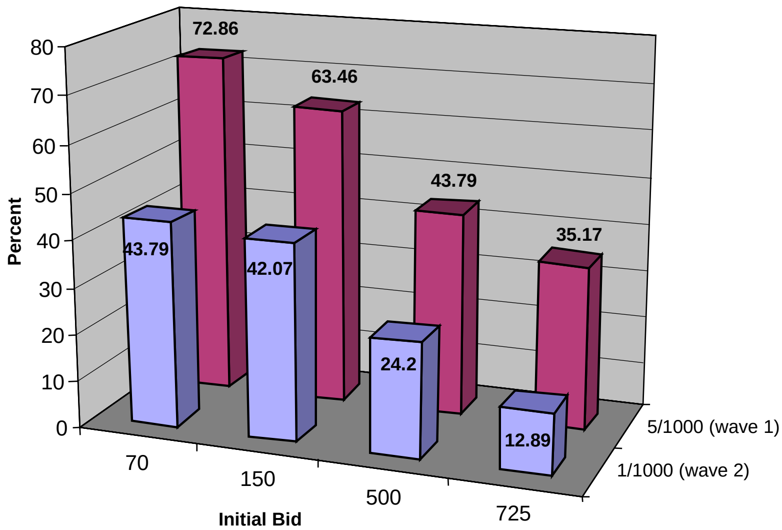

Fig. 2. Percent of "yes" responses by bid value: US Study.

dichotomous choice questions were asked if they would pay anything at all for the product. Those answering "yes" to both questions were asked to express their maximum WTP.[18] We elect not to use any of the open-ended WTP amounts in the analyses reported here. While positive WTP amounts reported on a continuous scale are easily accommodated within a maximum likelihood

[18] In Canada, 19.5% and 36.8% of the respondents indicated that they were not willing to pay anything at all for the 5 in 1000 and 1 in 1000 risk reductions, respectively. The corresponding US figures are 22.0% and 37.7%.

Table 6
Mean and median WTP for reduced mortality risk (US $[a])

	Canada		US	
	Median WTP	Mean WTP	Median WTP	Mean WTP
5 in 1000 risk reduction (wave 1)	253	466	350	770
	(17.1)	(33.6)	(28.7)	(86.9)
	$n = 616$	$n = 616$	$n = 556$	$n = 556$
1 in 1000 risk reduction (wave 2)	131	370	111	483
	(18.2)	(48.6)	(14.0)	(74.0)
	$n = 292$	$n = 292$	$n = 548$	$n = 548$
Are the WTP figures for risk reductions of different sizes…				
Significantly different?	Yes	Yes	Yes	Yes
Wald test	23.74	2.65	56.10	6.32
p-value	<0.0001	0.10	<0.0001	0.01
Proportional to the size of the risk reduction?	No (ratio = 1.9)	No (ratio = 1.3)	No (ratio = 3.2)	No (ratio = 1.6)
Wald Test	19.0	32.0	7.3	18.7
p-value	<0.0001	<0.0001	0.007	<0.0001

Notes: Standard errors in parentheses. Estimates based on interval-data Weibull estimation and cleaned data (respondents with FLAG1 = 1 deleted). No regressors are included in the Weibull models. Sample from the US excludes people older than 80 years of age.

[a] In US dollars, using purchasing power parity of C$1.25 per US$1.

estimation framework, we were dissatisfied with the performance of the models we devised to accommodate zero WTP responses.[19]

For these reasons we use only the responses to the two rounds of dichotomous choice questions. Our statistical model is an interval-data model based on the Weibull distribution, and is estimated using the method of maximum likelihood. The log likelihood function of the responses is

$$\log L = \sum_i \log[F(WTP_i^U; \theta, \sigma) - F(WTP_i^L; \theta, \sigma)], \tag{7}$$

where $F(\bullet; \theta, \sigma)$ is the cdf of the Weibull distribution with shape parameter θ and scale σ ($F(y; \theta, \sigma) = 1 - \exp(-(y/\sigma)^\theta)$), and WTP^L and WTP^U are the lower and upper bounds of the interval around the respondent's WTP amount.[20] Table 6 reports estimates of median and mean

[19] We have attempted two such models. In the first, responses are treated as if they came from a discrete mixture with two components. The first component is a degenerate distribution where all respondents hold a value of zero for the risk reduction, while the second component is a well-behaved random variable (a Weibull) that takes only positive, real values. Unfortunately, the estimation routine experienced convergence problems for all but the simplest specifications. The second model is an interval/continuous data version of the "spike" model described by Kriström [10]. The implicit restriction imposed in this model that the coefficients be the same for all respondents regardless if they hold zero or positive WTP amounts appears to be violated in our data, at least for the US study.

[20] Some authors have questioned the appropriateness of combining responses to the initial and follow-up payment questions, as is often done in double-bounded models of WTP, arguing that there is empirical evidence of a systematic shift in WTP between payment questions [1,8]. To explore this matter, we estimated single-bounded models of WTP, which rely on responses to the initial payment questions and ignore the follow-ups. Unfortunately, the (unconditional)

annual WTP based on Eq. (7) without using covariates.[21,22] We exclude from our models those respondents who failed both the initial probability test and probability choice questions described above, whether or not they subsequently corrected their answers (FLAG1 = 1). We also exclude respondents over the age of 80.[23]

In both Canadian and US samples, mean and median WTP are significantly larger for a 5 in 1000 risk reduction than for a 1 in 1000 risk reduction; however, neither mean nor median WTP increases *in proportion* to the risk reduction size. Still, median WTP does show more sensitivity to risk reduction size than mean WTP: in both samples the ratio of median WTP for a 5 in 1000 risk reduction to median WTP for a 1 in 1000 risk reduction is larger than the corresponding ratio of mean WTPs.[24]

(*footnote continued*)
estimates of mean WTP that result have large standard errors, making the comparison with WTP figures from double-bounded models inconclusive. The median WTP figures from single-bounded models have more reasonable standard errors, and are higher than double-bounded estimates for Canada (435 C$ vs. 394 C$, respectively), but lower than the double-bounded figures for the US (223 vs. 346 US$ respectively).

It is important to note that the marginal effect on WTP of individual characteristics such as age, income, gender, race and education does not change appreciably when we move from single-bounded to double-bounded models of WTP. This is true for both the US and Canada. For example, when we run the single-bounded version of regression (2) of Table 9 using the data from the US study, the coefficients on the age dummies are estimated to be 0.17, −0.33, and −0.41, respectively. The coefficients on the MALE and BLACK dummies are −0.55 and 0.75, respectively, and the coefficients on education and income per household member (in thousand dollars) are equal to −0.13 and 0.02. We therefore conclude that the coefficients on the age group dummies and education are insignificant, as in the double-bounded models, but those on MALE, BLACK, and the income variable are strongly significant, and are very similar to the corresponding coefficients from the double-bounded model. The intercept is higher for the single-bounded than for the double-bounded model. In sum, this suggests that opting for the double-bounded approach does not alter the marginal effect of the covariates on WTP, at least in the US study.

Finally, we experimented with the bivariate model [2], but it resulted in large, implausible mean WTP figures. This finding, along with the observation that double-bounded models do not significantly alter estimates of the marginal effects of covariates on WTP and afford large gains in efficiency, prompted us to use the double-bounded approach throughout the paper.

[21] All figures are expressed in US dollars, based on a purchasing power parity of C1.25 per US$1.

[22] Mean WTP is computed as $\hat{\sigma} \cdot \Gamma(1/\hat{\theta} + 1)$, and median WTP is $\hat{\sigma} \cdot (-\ln 0.5)^{1/\hat{\theta}}$ where the hats denote estimates. To compute standard errors around mean WTP, we drew samples of 20,000 observations from a multivariate normal distribution centered on the estimated Weibull parameters with covariate matrix equal to the inverse of the information matrix of log likelihood (7). We used the values drawn from this distribution to compute 20,000 estimates of mean WTP. The standard deviation of this vector of artificially generated mean WTP values is the standard error of the estimate of mean WTP shown in Table 7. For median WTP, this simulation-based approach is compared with the delta method. The two approaches produce very close standard errors around median WTP.

[23] We decided to exclude respondents older than 80 years of age. In our preliminary analyses of the relationship between WTP and age, we found that in the oldest age bracket (age 80 and older) the relationship appeared to be non-monotonic and driven by a few responses implying relatively large WTP amounts. Since there were only 11 subjects in this age group, we deemed it safer to exclude these respondents from our analysis.

[24] These findings also appear when we pool the WTP responses for the 5 in 1000 risk reduction from wave 1 with the WTP responses for the 1 in 1000 risk reduction from wave 2 for each country. We use each set of pooled data to fit a Weibull model where the scale parameter, σ, is expressed as exp(intercept + WAVE1$^{*}\delta$), WAVE1 being a dummy variable that denotes whether the WTP responses refer to the 5 in 1000 risk reduction. The shape parameter is common across the two risk reductions. We find that WTP for the 5 in 1000 risk reduction is 2.5 times WTP for the 1 in 1000 risk reduction in the US sample, and 1.92 times in the Canada sample, which rejects the hypothesis of proportionality.

Table 7
Implied estimates of the value of a statistical life (US $)

Magnitude of risk change	Canada		US	
	From median WTP	From mean WTP	From median WTP	From mean WTP
5 in 1000	506,000	933,000	700,000	1,540,000
1 in 1000	1,312,000	3,704,000	1,110,000	4,830,000

Note: VSLs computed using annual risk changes, i.e. 5 in 10,000 and 1 in 10,000.

The corresponding estimates of VSL are shown in Table 7. These are calculated using the mean annual WTP (from Table 6), and dividing it by the *annual* risk reduction implied by the product/action described to the respondent.[25] Given the similarity of the WTP figures from the Canada and US studies, it is not surprising that the VSLs are also very similar across the two studies.

The VSL estimates for Canada range from $506,000 and $933,000, when computed using WTP for the 5 in 1000 risk reduction, compared to $700,000 and $1.54 million for the US. When based on mean WTP for the 1 in 1000 risk reduction, the VSL estimates increase, reaching upwards of ($3.7 million for Canada and $4.8 million for the US. Because WTP is generally not proportional to the size of the risk change, VSL estimates are larger when calculated using WTP for the smaller risk change. Even the more generous VSL amounts are well below the standard estimate used by USEPA in its major peer-reviewed study of the costs and benefits of the Clean Air Act ($6.1 million, 1999$) [20] and in all recent EPA regulatory impact analyses. We do find, however, that these more generous estimates are close to the average VSLs from the five stated preference studies included in the set of studies on which the EPA default VSL is based ($3.7 million (1999$)) [20, Table H–1].

Returning to the issue of sensitivity to scope, when we distinguish respondents by the degree of confidence they have in their answers, median WTP does, indeed, increase in proportion to the size of the risk reduction. This is shown in Table 8. After each WTP question, we asked respondents to indicate the level of certainty they had in their responses on a scale from 1 to 7. In Table 8, certainty levels of 6 and 7 are categorized as ‘‘More Confident’’ and all other certainty levels as ‘‘Less Confident.’’ When attention is restricted to more confident respondents, the ratio between the median WTP amounts is 3.3 for Canada and 8.9 for the US, with neither figure statistically different from 5. Although we cannot reject the null hypothesis of strict proportionality for median WTP for more confident respondents, we warn the reader that these results should be interpreted with caution, because of the large confidence intervals around the point estimates of the ratio between WTP for the two different risk reductions.

(*footnote continued*)
Allowing the scale parameter of the Weibull to depend on additional variables, such as age dummies, gender, race, income and family and respondent health status dummies does not alter this conclusion.

[25] Assuming that the respondents spread the risk reduction evenly over the 10 years, this approach overcomes the difficulty of having to choose a discount rate for the 10 annual payments.

Table 8
The effect of confidence on median WTP for reductions to mortality risk (US$)

Magnitude of risk reduction	Canada median WTP		US median WTP	
	More confident	Less confident	More confident	Less confident
5 in 1000 (wave 1)	414	268	205	445
	(48.1)	(36.2)	(41.8)	(36.1)
	$n = 267$	$n = 349$	$n = 244$	$n = 311$
1 in 1000 (wave 2)	126	136	23	236
	(29.4)	(22.9)	(9.4)	(25.3)
	$n = 139$	$n = 151$	$n = 298$	$n = 250$
Are the WTP figures for risk reductions of different sizes…				
Significantly different?	Yes	Yes	Yes	Yes
Wald test	25.95	9.51	18.04	22.48
p-value	<0.0001	0.002	<0.0001	<0.0001
Proportional to the size of the risk reduction?	Yes (ratio = 3.3)	No (ratio = 2.0)	Yes (ratio = 8.9)	No (ratio = 1.9)
Wald test	1.99	11.8	2.04	31.21
p-value	0.15	0.0006	0.15	<0.0001

Note: Standard errors in parentheses. Estimates based on cleaned data (respondents with FLAG1 = 1 deleted; respondents older than 80 excluded).

5.3. The impact of age and health on WTP

To examine the impact of age and health on WTP, we add covariates to the Weibull model.[26] The life-cycle model presented in Section 2 suggests that current wealth and a person's future income stream should influence his WTP for a reduction in risk of death. So should covariates that may alter the individual's estimate of his chances of surviving the next 10 years, $(1 - D_j)$. These include, in addition to a person's own age and health, family health history, race and gender. These same variables, of course, should influence estimates of future survival probabilities, so their net impact on WTP is uncertain.

We present results of interval-data Weibull regressions of WTP in Table 9 for the 5 in 1000 risk reduction based on wave 1 for both the US and Canadian samples. The Weibull regressions assume that $WTP = \exp(x_i\beta)WTP_0^{1/\theta}$, where WTP_0 is WTP when all covariates are set to zero, which is distributed as a Weibull with shape parameter θ and scale 1. A log transformation produces the equation $\log WTP = \mathbf{x}_1\beta + (1/\theta)\varepsilon$, where $\varepsilon = \log WTP_0$ follows the type I extreme value distribution.

The first model presented in Table 9 examines the effects of age on WTP using dummy variables to represent age groups 50–59, 60–69, and 70 and older. When these dummies are included in the

[26] We also computed the percentage of "yes" responses to the initial payment question by age group. In Canada, 54.6% of the 40–49 year-olds were willing to pay the initial bid for the 5 in 1000 risk reduction. The percentage of "yes" responses was 52.6 among 50–59 year-olds, 50.4 among the 60–69 year-olds, and 50.0 for those respondents older than 70 years of age. In the US, the corresponding percentages are 52.6, 54.8, 54.2, and 46.7. The chi square tests indicate no significant differences across the two studies in the percentage of "yes" responses for each group.

Table 9
Weibull interval-data regression results for 5 in 1000 risk reduction (Wave 1) (standard errors in parentheses)

Variable	Canada (Can $)			US (US $)			Pooled Canada-US (in US$)		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
Intercept (ages 40–49)	6.42**	6.98**	6.81**	6.39**	6.95**	6.53**	6.36**	6.85**	6.65**
	(0.10)	(0.36)	(0.38)	(0.13)	(0.50)	(0.54)	(0.09)	(0.29)	(0.31)
Ages 50–59	0.15	0.19	0.17	0.05	0.13	0.071	0.11	0.13	0.09
	(0.16)	(0.16)	(0.16)	(0.21)	(0.23)	(0.23)	(0.13)	(0.13)	(0.13)
Ages 60–69	0.13	0.14	0.17	0.03	−0.07	−0.23	0.09	0.10	0.04
	(0.17)	(0.17)	(0.17)	(0.22)	(0.23)	(0.23)	(0.13)	(0.14)	(0.14)
Ages 70 and older	−0.34*	−0.35*	−0.33	−0.21	−0.17	−0.23	−0.27	−0.26*	−0.29*
	(0.20)	(0.21)	(0.21)	(0.23)	(0.25)	(0.26)	(0.15)	(0.15)	(0.16)
Male	—	−0.21*	−0.19	—	−0.43**	−0.45**	—	−0.31**	−0.32**
		(0.13)	(0.13)		(0.17)	(0.17)		(0.10)	(0.10)
Black	—	—	—	—	0.73**	0.71**	—	0.41	0.47
					(0.36)	(0.36)		(0.29)	(0.30)
Bottom 25% of distribution of income[a]	—	−0.23	−0.23	—	—	—	—	−0.20*	−0.21*
		(0.15)	(0.16)					(0.11)	(0.12)
Income per person	—	—	—	—	0.000013**	0.000016**	—	—	—
					(6.39×10^{-6})	(6.58×10^{-6})			
Education (years of schooling)	—	−0.03	−0.03	—	−0.05	−0.05	—	−0.02	−0.03
		(0.02)	(0.02)		(0.04)	(0.04)		(0.02)	(0.02)
Family history of chronic illness (excluding cancer)	—	—	0.26*	—	—	0.37*	—	—	0.29**
			(0.16)			(0.22)			(0.12)
Family history of cancer	—	—	−0.06	—	—	0.02	—	—	−0.01
			(0.13)			(0.17)			(0.10)
ER visit in last five years or hospitalization in last year	—	—	−0.07	—	—	0.63**	—	—	0.23
			(0.19)			(0.28)			(0.16)
No insurance	—	—	−0.03	—	—	0.29	—	—	−0.03
			(0.15)			(0.28)			(0.13)
Canada dummy	—	—	—	—	—	—	−0.16	−0.12	−0.11
							(0.10)	(0.10)	(0.11)
Scale parameter	1.27	1.27	1.26	1.44	1.41	1.37	1.34	1.34	1.32
	(0.07)	(0.07)	(0.07)	(0.09)	(0.09)	(0.09)	(0.05)	(0.05)	(0.05)
Sample size	616	605	605	551	474	474	1167	1079	1079

**Significance at the 5% level.
*Significance at the 10% level. FLAG1 = 1 deleted. US sample only includes people of ages no greater than 80 years.
[a] Bottom 25% of the distribution of household income is C$24,500 in the Canada sample, US$32,500 in the US sample.

Weibull model, the coefficients of the age 50–59 and 60–69 age group dummies are indistinguishable from the coefficient of the 40–49 age bracket (captured by the intercept). The coefficient of the oldest age bracket, however, is significantly lower for the Canada sample, at the 10% level. This remains true when other covariates are included in the regression. No age effects are found using the US sample.

Model 2 examines, in addition to age, the impacts of race, gender, income, and education on WTP. In the US study, income is measured as household income divided by household size. Information about the size of the household is not available for the Canadian sample. Instead, the Weibull regression includes a dummy that takes on a value of one if the Canadian respondent's household income is in the first quartile. In both samples WTP increases with income, although this effect is statistically significant only in the US.[27] In the US, blacks are willing to pay more for a 5 in 1000 risk reduction than whites and Hispanics, possibly due to their higher baseline risks, whereas males are willing to pay less. Education does not have a statistically significant effect on WTP.[28,29]

Model 3 examines the impact of family health history, health insurance and hospital admissions on WTP. These variables are insignificant in both samples, with two exceptions. Family history of chronic heart or lung disease increases WTP by 26% in Canada, and by 37% in the US. Specifically, respondents who had been admitted to the hospital for a heart or lung condition in the last year or who had been admitted to an Emergency Room for one of these conditions in the last 5 years are willing to pay 63% more to reduce their risk of death than persons who had not had such hospital visits.[30,31]

These results are robust to changing the criteria for "cleaning" the sample. For example, when we only exclude from the sample those respondents who confirmed the "wrong" answers to the probability test or to the probability *choice* questions (10 respondents for the Canada sample, and 22 for the US sample), we find that the coefficients on the age dummies, health status, education,

[27] The income elasticities of WTP implied by the coefficients of income per family member in the US sample are 0.26 (specification 2) and 0.33 (specification 3).

[28] Because baseline risk varies systematically with the respondent's age, gender and—in the US—race, it is not possible to identify the impact of baseline risk on WTP. Baseline risk was found to be insignificant, both for Canada and the US in Weibull regressions where it was the only determinant of WTP. Regressions with baseline risk and an interaction term between baseline risk and the respondent's acceptance of the risk figure we presented to them in the survey yielded coefficients implausible in both sign and magnitude.

[29] To further explore the impacts of age, gender, and health status of the respondent and of the respondent's family, for each country we pooled the WTP responses for the 5 in 1000 risk reduction from wave 1 and for the 1 in 1000 risk reduction from wave 2. As explained in footnote 18, we ran separate WTP regressions for each country, making sure to include a dummy for the size of the risk reduction. While in the US sample WTP is significantly lower for 60–69 year-olds (estimated coefficient −0.39, s.e. 0.19) and 70-year-olds and older (estimated coefficient −0.45, s.e. 0.23), in the Canada study it no longer varied significantly across age groups.

[30] The effects of age, race, gender, income and family health history are, in general, robust to the inclusion of variables indicating whether the respondent questioned baseline risk figures or the assumptions of the WTP scenario. (Table A.1, available at http://www.aere.org/journal/index.html, reports the results of adding the debriefing variables to Model 3). When significant, the coefficients on the debriefing variables are of the correct sign.

[31] We checked whether recent visits to the Emergency Room or hospitalizations implied that the respondent does not believe the risks stated to him in the survey, and thinks, instead, that his true baseline risk is higher. However, we found no significant difference in terms of acceptance of the baseline risks across those respondents who did and did not report having recently been to the Emergency Room or hospital.

income, and gender are virtually unchanged. Only the coefficient on the race indicator for the US sample is about 20% lower when the latter data-cleaning criterion is employed.

One question is whether individual characteristics, family health history, health insurance coverage and past hospital admissions for heart and respiratory disease affect WTP to the same extent in both samples. To answer this question, we pool the WTP responses for the 5 in 1000 risk reduction in the two countries, convert all values to US dollars, and include a country dummy in the WTP regressions.

Results are reported in Table 9 for all three specifications. Comparison across the specifications suggests that the oldest respondents hold WTP values that are 20–25% less than younger respondents, an effect that is significant at the 10% level or better.[32] Male respondents continue to be willing to pay less than females, but the WTP values of blacks are no longer statistically different from those of other respondents. Also, respondents with incomes in the bottom 25% of the income distribution in their respective samples report considerably lower WTP amounts than households with higher incomes.

As before, all else the same, having family members with a history of chronic heart or respiratory illness increases WTP by 34%, whereas insurance coverage, a history of family cancer and past hospital admission do not have a statistically discernible effect. The coefficient on the Canada dummy implies that Canadian respondents hold WTP values that are slightly smaller than those of US respondents of similar individual characteristics, but the difference between the two groups is not statistically significant.

That the two samples are similar in terms of the determinants of WTP is not limited to this additive specification. We performed likelihood ratio tests of the null hypothesis that *all* parameters of models 2 and 3 are the same across the two studies, and found that we could not reject the null hypothesis in either case.[33] In addition to recording hospital admissions, we control for respondent health status with a series of dummy variables indicating whether the respondent has ever been diagnosed with cardiovascular disease, chronic lung disease, high blood pressure or cancer.[34] Because these variables are correlated we add them one at a time.[35] Their coefficients appear in Table 10. In the Canada study none of the chronic health dummies is ever statistically significant. In the US study while the dummies for cancer and cardiovascular disease are never statistically significant, having high blood pressure raises WTP about one-third, even after controlling for family health history and hospital visits. Having a chronic respiratory illness significantly raises WTP when the latter variables are omitted from the equation, but not when

[32] We remind the reader that respondents were randomly assigned bid amounts. The results from the Weibull regressions of WTP on the age group dummies are confirmed by the patterns of "yes" responses to the initial payment question. To illustrate, we pooled the US and Canada data, and subsequently split this pooled sample into the group of respondents who were younger than, and older than, 70 years of age, respectively. In both groups, the percentage of "yes" responses declines with the bid, ranging from a high of 73% to a low of 32% among the younger group, and from 69% to 26% in the older group. With the second lowest bid as the only exception, the proportions of "yes" responses are lower in the older group at each bid value used.

[33] The race dummy is not included in the performance of these tests, since there are no blacks in the Canada sample.

[34] We have chosen to focus on indicators of chronic illness rather than indicators of functional limitation, as measured by the indices from SF-36. The latter are often correlated with the former; however, it is the former that are used more often in epidemiological studies. The latter are almost always insignificant in our regressions.

[35] We tried interacting each health dummy with the ER visit/Hospitalization variable. These were never significant.

Table 10
Effect of adding health variables one at a time to previous models (standard error in parentheses)

Characteristic	Canada		US		Pooled Canada and US	
	Model 2+	Model 3+	Model 2+	Model 3+	Model 2+	Model 3+
CARDIO[a]	0.16	0.19	0.19	−0.02	0.22	0.14
	(0.22)	(0.23)	(0.22)	(0.25)	(0.21)	(0.22)
Chronic respiratory illness (asthma, emphysema,	0.03	0.03	0.45**	0.29	0.16	0.11
chronic bronchitis)	(0.16)	(0.16)	(0.23)	(0.23)	(0.13)	(0.13)
High blood pressure	0.14	0.11	0.38**	0.35*	0.22*	0.16
	(0.16)	(0.17)	(0.19)	(0.19)	(0.12)	(0.12)
Cancer	0.51	0.53	0.28	0.19	0.34	0.30
	(0.36)	(0.36)	(0.31)	(0.31)	(0.21)	(0.21)

**Significance at the 5% level.
*Significance at the 10% level. FLAG1 = 1 deleted. US sample only includes people of ages no greater than 80 years.
[a] Definition of CARDIO: respondent has one or more of the following: angina pectoris, coronary disease, other heart diseases, and/or has had a myocardial infarction.

they are added. We conclude that having a chronic condition does not reduce WTP for mortality risk reductions and may even increase it.[36] This conclusion holds when we pool the data, as shown in Table 10, although the coefficients on the chronic condition dummies are almost always insignificant.

6. Conclusions and implications for policy

In this paper we have used contingent valuation to examine the effects of current age and health status on respondents' WTP for a product that would reduce their risk of death over the next 10 years. Economic theory is ambiguous about the impact of these variables on WTP to reduce risk of death. In general, WTP should be higher the lower an individual's chances of surviving the current period. It should also be higher, the greater the present discounted value of lifetime utility conditional on surviving the current period. In a comparative static sense, being older and having a chronic heart or lung condition both reduce an individual's chances of surviving the current period and, for this reason, tend to raise WTP. Older individuals and persons with chronic diseases, however, have fewer expected life years to look forward to, assuming they survive the current period. This may lower their WTP to reduce current risk of death, depending on how the present value of utility of consumption changes with age.

Our results suggest that, for health status, the former effect balances out the latter and may even dominate it: Persons with chronic heart and lung disease are willing to pay at least as much to reduce their risk of dying as persons without these diseases. In our US sample, WTP is significantly greater (holding age, gender and income constant) for persons with high blood pressure than for those without. There are no statistically significant differences between the two groups in Canada. WTP remains significantly higher for persons with high blood pressure when data are pooled across countries. Chronic respiratory illness and heart disease have no statistically significant effect on WTP in either country.

As regards age, we find weak support for an age effect, and only for respondents 70 years old and above. Respondents in our Canada sample over age 70 were willing to pay about one-third

[36] Tables 9 and 10, clearly use a reduced-form model relating WTP with age and health status. For the US data, we estimated a structural model based on two simultaneous equations: (i) one relating remaining life years (subjectively assessed by the respondent in the survey) to individual characteristics, including age, health status, income, education, gender and race, and (ii) one relating WTP to age, health status, income, education, gender and race, *and* remaining lifetime. The two equations are simultaneous because it is likely that unobserved individual characteristics influence both self-assessed remaining life years and WTP. To ensure identification, Eq. (i) includes a dummy for whether the respondent's mother is still alive, how old she was when she died (if no longer living), and mental health score—a measure of psychological distress, or lack thereof—from the SF-36 questions. Briefly, we found that age and health influenced remaining life years, but have no direct effect on WTP. As in the reduced form reported in Table 9, WTP for the US sample in the structural system is influenced by income, gender, race, and family chronic health history. The coefficient of remaining life years, however, was negative and insignificant, which is against the expectations from the theoretical model, suggesting that Eqs. (i) and (ii) make up a system of seemingly unrelated, rather than simultaneous, equations. Perhaps we should interpret Eqs. (i) and (ii) as being derived from the life-cycle model of WTP for a risk reduction under rather restrictive assumptions, namely, that $u'(X)/u(X)$ is constant with respect to age, and that the discount rate is zero.

less than their younger counterparts to reduce their risk of dying by 5 in 1000 over the next 10 years. In the US, persons over age 70 were willing to pay about 20% less than persons below that age for a 5 in 1000 risk reduction; however, this effect was not statistically significant. Pooling the data between countries for the 5 in 1000 risk reduction resulted in a 25% reduction in WTP over age 70 that was statistically significant.

Our results do not necessarily support current practices in benefit–cost analyses at the USEPA and elsewhere. The standard practice of using a common VSL for valuing risk reductions irrespective of the age and health status of those benefiting seems questionable for the oldest age groups. The growing practice of apportioning VSLs over life expectancy—the VSLY approach—is not supported by our results. Yet, our results do support Health Canada's approach of using age-adjusted VSL estimates in its economic assessments, applying a VSL of C$5 million (or US$4 million) to exposed populations under 65 years of age and using an adjustment factor of 0.75 for populations aged 65 years and over [3].[37]

Furthermore, our results do not necessarily endorse the specific VSLs used by agencies. For instance, the USEPA currently uses a central VSL estimate, based primarily on labor market studies, equal to about $6 million (1999 US$) for all ages [20]. Our estimated VSLs are only from one-sixth (relying on the results for a 5/10,000 annual risk change) to three-quarters (relying on the results for the 1/10,000 risk change) of this value. Finally, our results call into question the use by some public health agencies of measures such as QALYs, which, by the way the index is constructed, devalue benefits to the infirm and reduce benefits proportionally to the age of the beneficiary.

References

[1] A. Alberini, B. Kanninen, R.T. Carson, Modeling response incentive effects in dichotomous choice contingent valuation data, Land Econ. 73 (3) (1997) 309–324.
[2] T.A. Cameron, J. Quiggin, Estimation using contingent valuation data from a 'dichotomous choice with follow-up' questionnaire, J. Environ. Econ. Manage. 27 (3) (1994) 218–234.
[3] Canada-Wide Standards Development Committee for PM and Ozone, Discussion Paper on Particulate Matter and Ozone, Canada-Wide Standards for Consultation, National Multi-Stakeholder Consultation Workshop, Calgary, Alberta, May 26–28, (1999).
[4] M. Cropper, F. Sussman, Valuing future risks to life, J. Environ. Econ. Manage. 19 (1990) 160–174.
[5] EOP Group, Inc., Life-Year Analysis of Premature Mortality Benefits in the December 1996 Particulate Matter Proposed NAAQS, 1997 (February).
[6] J. Fernandez-Villaverde, D. Krueger, Consumption Over the Life-Cycle: Some Facts from Consumer Expenditure Survey Data, NBER Working Paper 9382, 2002 (December).
[7] M.R. Gold, L.B. Russell, J.E. Siegel, M.C. Weinstein (Eds.), Cost-Effectiveness in Health and Medicine, Oxford University Press, New York, 1996.
[8] J.A. Herriges, J.F. Shogren, Starting point bias in dichotomous choice valuation with follow-up questioning, J. Environ. Econ. Manage. 30 (1) (1996) 112–131.
[9] M.W. Jones-Lee, M. Hammerton, P.R. Phillips, The value of safety: results of a national sample survey, Econ. J. 95 (1985) 49–72.
[10] B. Kriström, Spike models in contingent valuation, Amer. J. Agr. Econ. 79 (1997) 1013–1023.

[37] These adjustments are based on Jones-Lee et al. [9].

[11] A. Krupnick, A. Alberini, M. Cropper, N. Simon, B. O'Brien, R. Goeree, M. Heintzelman, Age, health and the willingness to pay for mortality risk reductions: a contingent valuation study of Ontario residents, J. Risk Uncertainty 24 (2002) 161–186.
[12] M.J. Moore, W.K. Viscusi, The quantity-adjusted value of life, Econ. Inq. 26 (1988) 369–388.
[13] K.H. Morales, L. Ryan, T.-L. Kuo, M.-M. Wu, C.-J. Chen, Risk of internal cancers from arsenic in drinking water, Environ. Health Perspect. 108 (2000) 655–661.
[14] C.A. Pope, et al., Particulate air pollution as a predictor of mortality in a prospective study of US adults, Am. J. Respir. Crit. Care Med. 151 (1995) 669–674.
[15] J. Schwartz, Particulate pollution and daily mortality in Detroit, Environ. Res. 56 (1991) 204–213.
[16] J. Schwartz, Air pollution and daily mortality in Birmingham, Alabama, Amer. J. Epidemiol. 137 (10) (1993) 1136–1147.
[17] J. Schwartz, D.W. Douglas, Increased mortality in Philadelphia associated with daily air pollution concentrations, Am. Rev. of Respir. Dis. 145 (1992) 600–604.
[18] D.S. Shepard, R.J. Zeckhauser, Life-cycle consumption and willingness to pay for increased survival, in: M.W. Jones-Lee (Ed.), The Value of Life and Safety, North Holland, Amsterdam, the Netherlands, 1982.
[19] C. Skrzycki, Under fire, EPA drops the 'senior death discount, Wash Post, (May 13, 2003) E01 (2003).
[20] US Environmental Protection Agency, The Benefits and Costs of the Clean Air Act, 1990–2010, Report to the US Congress, 1999 (November).
[21] J.E. Ware Jr., M. Kosinski, S. Keller, SF-36 Physical and Mental Health Summary Scales A User's Manual, Quality Metric, Lincoln RI, 1997.
[22] M.E. Yaari, Uncertain lifetime, life insurance and the theory of the consumer, Rev. Econ. Stud. 32 (1965) 137–150.