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COMPLEMENTARITY AND THE MEASUREMENT OF INDIVIDUAL RISK TRADEOFFS: ACCOUNTING FOR QUANTITY AND QUALITY OF LIFE EFFECTS

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ABSTRACT

This paper considers the factors responsible for differences with age in estimates of the wage compensation an individual requires to accept increased occupational fatality risk. We derive a relationship between the value of a statistical life (VSL) and the degree of complementarity between consumption and labor supplied when health status serves as a potential source of variation in this relationship. Our empirical analysis finds that variations in an individual's health status or quality of life and anticipated longevity threats lead to significant differences in the estimated wage/risk tradeoffs. We describe how extensions to the specification of hedonic wage models, including measures for quality of life and anticipated longevity threats, help to explain the diversity in past studies examining how the estimated wage–risk tradeoff changes with age.

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I. Introduction

Recent research has identified how complementarity relationships help in resolving some empirical puzzles in economics. In the first of these papers, Chetty [2006] develops an upper bound on the coefficient of relative risk aversion using labor supply elasticities. His relationship implies that the bound varies with the degree of complementarity between consumption and labor supplied. A second paper by Hall and Jones [2007] uses a form of complementarity to explain the rationality of an increasing income share for health expenditures as incomes grow. In particular, Hall and Jones assume consumption goods and health services are weakly complementary. We argue that combining these two independent insights provides a possible explanation for yet another puzzle largely associated with microeconomic applications—the relationship between age and the marginal willingness to pay for risk reductions (the value of a statistical life or VSL) and the diverse empirical findings from revealed preference studies designed to examine this relationship.[1]

We use Chetty's result to show that complementarity between consumption and labor supply implies, in the absence of other considerations, a lower estimate for the value of a statistical life than would be implied by models that assume independence. We hypothesize that health status may well be a third influence on individual choice. That is,

[1] Chetty describes the logic for our proposed analysis as follows: "If complementarity between consumption and labor is sufficiently strong, even highly risk averse individuals may choose not to reduce labor supply when wages rise because increased consumption makes work less painful" (p.1821).

a person’s state of health may influence the extent of complementarity between labor supply and consumption.

This logic is consistent in broad terms with the arguments in Hall and Jones [2007] where health expenditures contribute to maintaining a health state. Improvements in health increase the value of leisure and change the types of goods consumed. Thus, improved health increases the demand for leisure and with it demands for a different mix of consumption goods, which in turn reduces the apparent degree of complementarity between labor supply and a fixed weight measure for aggregate consumption. Thus, differences in the degree of complementarity between labor supply and consumption among individuals with different health states offer one potential explanation for why we might observe differences in estimated VSLs across individuals of different ages.

Finally, identifying the relationships among health status, complementarity, and estimated risk/wage tradeoffs helps to reconcile diverse findings regarding the relationship between age and the VSL in the revealed preference literature. Viscusi and Aldy [2007] suggest lower VSL estimates among the older workers in their sample, with a sample mean age of 39; Smith et al. [2004] fail to find lower VSL estimates for the older workers in their sample, with a sample mean age of 55. Neither analysis, however, controls for differences in health status among workers. If the distribution of health status varies among those workers in the two samples, then this variation could help to explain the observed differences in empirical results.

This paper examines the effects of an individual’s health status or quality of life as well as a person’s latent perceptions of his (or her) longevity on estimates of the VSL. Our analysis considers a sample of older workers who participated in the Health and

Retirement Study (HRS).[2] We find these measures for quality and quantity remaining of life influence the risk/wealth tradeoffs in ways that are consistent with our complementarity explanation. Specifically, we exploit the panel dimension of the HRS to define a measure of subjective quality of life based on a survey response given two years before labor market choices are reported. The motivation for our measure of the (subjective) quantity of life remaining stems from results in Smith et al. [2001] that suggest a respondent's perceived longevity can be viewed as a latent variable from the analyst's perspective. While this latent variable influences the other risks an individual will assume, it is not completely reflected in subjective longevity assessments elicited at the time the labor market behavior is recorded. Thus, we use a future death (or another serious health condition) as an instrument for a respondent's current perceptions of personal longevity or as a subjective measure of the "quantity" of life remaining. Typically, the analyst is unable to take into account how these factors would affect risk taking behaviors; the HRS panel structure allows us to consider them.

Section two provides a brief summary of the policy context for the questions associated with the sources of heterogeneity in risk preferences. The third section describes the theoretical logic that underlies the tests in our basic empirical model. Section four summarizes our data and findings and the last section discusses their implications.

II. Policy Context

In 2002, EPA's Regulatory Impact Analysis (RIA) for the Clear Skies initiative (U.S. Environmental Protection Agency [2002]) included an adjustment to the VSL

[2] The HRS (Health and Retirement Study) is sponsored by the National Institute of Aging (grant number NIA U01AG09740) and conducted by the University of Michigan. Where possible, we rely on the RAND

estimates used to monetize the benefits from reduced risks experienced by older populations. Specifically, the RIA included an estimate of the mortality benefits of the proposed regulations using a VSL estimate discounted by approximately thirty-six percent. The resulting controversy over this "senior discount" has affected subsequent policy guidance, scientific reviews of the literature, and research. For example, following this controversy the Office of Management and Budget [2003] has suggested that regulatory agencies consider the possibility of using cost effectiveness analysis and quality adjusted life years to acknowledge that policies yield risk reductions to heterogeneous populations. In a recent advisory letter to the EPA Administrator, the Environmental Economics Advisory Committee of EPA's Science Advisory Board (SAB) addressed the policy use of VSL estimates adjusted for differences in life expectancy. The letter concluded that:

> "…The VSL may increase, decrease or remain constant as life expectancy decreases. The relationship between the VSL and life expectancy is therefore an empirical matter. In practice, because life expectancy is difficult to observe, the Agency will have to relate the VSL to factors related to life expectancy – namely age and health status. <u>Although the literature on the relationship between age and VSL is growing, the Committee does not believe that it is sufficiently robust to allow the Agency to use a VSL that varies with age</u>"
> (Morgan and Cropper [2007], emphasis added).

This conclusion is important because there is a sharp contrast between summaries of the results from revealed preference studies that rely on evidence from the labor market and the estimates from stated preference studies.[3] In a review of the revealed preference research that has examined the age-VSL relationship, Aldy and Viscusi [2007b] conclude that the VSL varies with age and that there is sufficient information to

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Corporation's cleaned version of the HRS. Otherwise, we use the raw HRS data.

[3] See also the review of the stated and revealed preference literature in Evans and Smith [2006].

suggest "…VSL increases with age, peaks in mid-life, and subsequently declines" (p. 257). By contrast, Krupnick [2007] concludes, based on his review of the stated preference findings, that "…considering the weight of the evidence, the implication is that for countries that apply a single VSL to adults of all ages, there is insufficient information and consensus to make a reasoned decision to switch to using either different VSLs for different ages (in a private-good context) or a VSLY [value of a statistical life year] which imposes a linear (discounted) relationship between life-years remaining and the VSL" (pp. 275-276, bracketed phrase added).

Even if the empirical issues could be resolved, a decision about how, in a normative sense, to treat heterogeneity in risk preferences elicits a broad range of emphatic (if not dramatic) responses. The same SAB advisory letter we cited earlier concludes its recommendations by noting that

> "…while there is no denying the reality of income effects, it is a policy judgment, not a scientific question, whether the same VSL should be employed in all regulatory decisions across a society or different values should be closer depending upon the preferences and income of the population affected by a specific regulation" (Morgan and Cropper [2007]).

Aldy and Viscusi [2007b] confirm the importance of these considerations using a new set of estimates of the wage/risk tradeoffs that provide age-specific VSLs. They compare the benefit estimates for the Clear Skies Initiative derived using their age- and industry-specific measures of the VSL to the estimates derived using the EPA's constant, consensus VSL estimate. Their results suggest about a forty percent reduction in the benefits of the Clear Skies Initiative for older populations using the age-differentiated VSL measures.

Of course, using a constant VSL to estimate the aggregate benefits of a policy that affects a population with different VSLs is inconsistent with the basic principles underlying benefit-cost analysis. As Sunstein [2004] observes, "a movement toward greater individuation of VSL would undoubtedly be extremely controversial. *But it is not merely consistent with that theory that justifies current practice; it is mandated by it. In principle, government should not force people to buy protection against statistical risks at a price that seems to them excessive*" (p. 5, emphasis added).

Sunstein's argument for complete individuation is unlikely to be politically or practically feasible. Most of the policy uses for estimates of income-risk tradeoffs are in short-term policy analyses with modest resources for new research and short time horizons for completion. Moreover, in many cases, the epidemiological analyses used to identify the relationship between environmental quality and health risks simply do not have the ability to carefully distinguish how each policy will impact sub-groups in a heterogeneous population. These limitations are noteworthy because they impose genuine practical constraints on what constitutes feasible policy analysis. In reality, the uncertainties in measurement largely trump the political debates about the fairness of a framework that allows a level of individuation that would be consistent with revealed preference measures of risk tradeoffs.

These concerns aside, the logic of the hedonic wage equilibrium suggests heterogeneous risk preferences and skills in adapting to job risk will cause workers to sort among jobs based on their personal attributes.[4] Both an individual's quality of life and

[4] The basic logic of the equilibrium described in a hedonic model has generally been recognized as a matching process (see Rosen [1974] for an early discussion). Ekeland, Heckman, and Nesheim [2002] provide a brief overview of how this interpretation of the market equilibrium affects the identification of structural parameters governing the behavioral functions in the process and what can be identified.

quantity of life remaining should affect how he or she makes choices in the labor market. As a result of this sorting, the market equilibrium will reveal different wage/risk tradeoffs across workers based on each individual's personal circumstances. The empirical challenge in measuring the role of each person's quality and quantity remaining of life for risk/compensation tradeoffs is to identify instruments that adequately reflect these dimensions and permit the analysis to overcome the effects of endogeneity of contemporaneous measures of these effects.

Given the potential constraints on the feasible responses to the empirical evidence from a policy perspective, why study the sources of differences in VSL? We believe research on the behavioral sources for differences in the tradeoffs people would make to reduce risk can influence other dimensions of risk assessments by identifying the important features of the heterogeneity in people's risk tradeoffs. That is, with this information it is possible for analysts to focus on a subset of key distinctions that can arise in selecting the components of the overall risk analysis for special emphasis. Inevitably the preparation of each regulatory impact analysis faces resource and time constraints. Nonetheless, even if the analyst never explicitly assigns different values to each group, knowledge of the potential for differences in the tradeoffs among different groups can contribute to the design of sensitivity analyses to gauge how policy judgments would change with individuation.

Our application does not have sufficient information to develop reliable measures of the VSL for specific groups of individuals. Instead, we use the results to estimate the wage-risk tradeoff for one group of individuals relative to another group.

Recently Shogren and Stamland [2002] focus on the information to be derived by describing how individuals' observable characteristics might influence what can be learned from equilibrium schedules.

III. Complementarity and the VSL

The motivation for our empirical analysis in the next section combines insights from three recent papers by Smith, Pattanayak, and Van Houtven [2003], Chetty [2006], and Hall and Jones [2007] on the relationships among the Arrow-Pratt coefficient of relative risk aversion, the compensated labor supply elasticity, the degree of complementarity between consumption and labor supply, and the value of a statistical life. The first (Smith et al. [2003]) develops a simple link between the uncompensated labor supply elasticity and the VSL for a specific preference specification, based on Burtless and Hausman's [1978] framework. In particular, Smith et al. find that the wage risk tradeoff can be expressed as in equation (1),

$$\frac{dw}{dp} = \frac{w}{(1-p)\varepsilon_{lw}} \tag{1}$$

where p represents fatality risk, w is the wage, and ε_{lw} denotes the uncompensated (Marshallian) labor supply elasticity. Chetty [2006] provides the second insight, which confirms a more general link between labor supply elasticities and estimates of the VSL. Using a general indirect utility function, denoted $v(c,l)$, defined in terms of Hicksian composite goods for the inter-temporal choices of consumption goods (c) and labor supply (l) over the life cycle, Chetty [2006] develops a relationship between the Arrow Pratt coefficient of relative risk aversion (R) and the ratio of the income elasticity of labor supply to the substitution elasticity of labor supply (or compensated labor supply elasticity), and the proportional response of consumption to changes in labor supply. Equation (2) summarizes the basic form of Chetty's result (his equation (7))[5]

$$R = -\frac{v_{cc}(c,l)}{v_c(c,l)}c = -\left(\frac{1}{1-\alpha}\right)\frac{\varepsilon_{lm}}{\varepsilon_{lw}^c} + \frac{1}{\alpha}\varepsilon_{ucl} \tag{2}$$

where $\alpha = \dfrac{w \cdot l}{m + w \cdot l}$ with *m* and *w* representing unearned income and the wage rate, respectively. Labor income is given by $w \cdot l$. ε_{lm}, ε_{lw}^c, and ε_{ucl} denote the income elasticity of labor supply, the compensated labor supply elasticity, and the elasticity of the marginal utility of consumption with respect to labor respectively.[6]

The compensated and Marshallian elasticities of labor supply can be related using a Hicks-Allen equation as in (3)

$$\varepsilon_{lw}^c = \varepsilon_{lw} - \frac{w \cdot l}{m}\varepsilon_{lm} \tag{3}$$

Solving equation (3) for ε_{lm} and substituting the result into (2), with some manipulation, we have:

$$\varepsilon_{lw} = \varepsilon_{lw}^c(1 + \varepsilon_{ucl} - \alpha R)\,. \tag{4}$$

Increased complementarity between consumption and labor supplied increases ε_{ucl}. For a given coefficient of relative risk aversion (*R*), share of wage income in total

[5] Equation (2) is also related to another connection more closely linked to environmental applications, the willingness to pay (WTP) and willingness to accept (WTA) disparity. It is an extended version of the relationship Hanemann [1991] used to explain dramatic differences between WTP and WTA for non-market environmental services by establishing a simple connection between the price flexibility of income for non-market goods, the income elasticity of demand for these goods, and an elasticity of substitution between these non-market environmental services and a composite of all other goods.

[6] The relationship between the Arrow-Pratt coefficient of relative risk aversion and the income elasticity of the VSL is closely related to our discussion here. This relationship was discussed by Eeckhoudt and Hammitt [2001] in an appendix to a paper that focused on the role of background risk for the properties of wage risk tradeoffs. More recently, Kaplow[2005] develops an analogous relationship to discuss the properties of the income elasticity of the VSL.

income (α), and compensated labor supply elasticity (ε^{c}_{lw}), increased complementarity implies a higher Marshallian labor supply elasticity (ε_{lw}) and, by equation (1), a lower VSL. If good health increases the demand for leisure and leisure- related consumption goods, then this change can in turn reduce the apparent degree of complementarity between labor supply and a consumption aggregate as depicted in the Chetty formulation.

Thus, to the extent improved health increases the complementarity between leisure and consumption we should expect an individual would require greater compensation to accept risk (a higher VSL). With poorer health (i.e., a poorer quality of life), leisure is less complementary to consumption, the uncompensated labor supply elasticity will be larger and the VSL smaller. Of course, it is important to acknowledge the other parameters included in equations (2) and (4) are not constants and can be expected to change with health status. Our argument merely provides motivation for one possible channel of influence on the labor supply elasticity – health status as an influence on the complementarity between consumption and labor supply. This connection in turn helps to explain how an individual's health status can affect her willingness to accept wage compensation to accept risk and thus measures of the VSL.

While health status can be considered a measure of the quality of life, our analysis cannot ignore the potential for the remaining quantity of life to impact the estimated wage risk tradeoffs (Pratt and Zeckhauser [1996]). Therefore, our empirical analysis contains a measure for health status (or quality of life) as well as an instrument intended to reflect each individual's unobserved subjective longevity expectations (or perceptions about the quantity of life remaining). Estimates from past studies (including our own) that have examined the relationship between age and the VSL will reflect the joint impacts of differences across individuals in the effects of health status and serious threats to longevity. To the extent health status, subjective beliefs about longevity, or some composite of the two effects are correlated with age and separate measures of these effects are not included in the specification of the wage equation, the omission has the potential to influence estimates of the effects of age on risk tradeoffs. Our analysis seeks to disentangle these effects. Unfortunately, we do not have sufficient information to allow a formal test of the hypothesis that differences in estimated VSLs with health status stem from differences in the extent of complementarity between consumption and leisure (labor supply).

IV. Empirical Model, Data, and Results

Our empirical analysis relies on the observed labor market experiences of respondents in the Health and Retirement Study (HRS). We exploit the panel dimension of this survey to distinguish measures of the increased risk of death and the deteriorating quality of life for this sample of older adults. In earlier research (Smith et al. [2004]) we did not take full advantage of these opportunities. Our earlier analysis was comparable to the rest of the literature in that we assumed difference in individuals' assessments of the quality of their lives as well as their beliefs about longevity would not independently influence risk tradeoffs revealed in the labor market choices; rather, we assumed age would proxy for these effects. We also used an average job risk measure for all respondents in an industry and assumed this aggregate measure was relevant for the largely older workers represented in the HRS.

Recently, Aldy and Viscusi [2003, 2007a, 2007b] have proposed the use of the Bureau of Labor Statistics' (BLS) Census of Fatal Occupational Injuries (CFOI) to develop a refined measure of job risk that accounts for the ages of workers. They structure mortality risk cells for each industry (by 2-digit SIC code) and age group. Our analysis uses the same on-the-job fatality risk measures, based on workers' ages and industries.

We consider three sets of models. The first is our primary model. The model controls for demographic attributes, occupation, a measure of risk tolerance (again as a qualitative variable), and the age- and industry-specific job risk. This model treats age non-parametrically, using a set of categorical variables that distinguish different age intervals. We consider interactions of job risk with age, risk tolerance, as well as the quality of life measure and our instrument for perceived longevity. To evaluate the sensitivity of our findings to the treatment of age, our second set of models includes age and age squared as an alternative description for how age has been hypothesized to influence wages along with risk tradeoffs. This specification also includes the same controls as in the primary model. The third set of models excludes age and risk tolerance measures to evaluate whether the effects of the health status variable and the longevity instrument are influenced by the presence of these other controls. Finally, the last set of models considers the effects of health status and longevity separately to evaluate whether these variables offer robust measures of the separate influences of individual effects of the quality and the quantity of life remaining. Rather than report all the estimates for each model we present the key parameter estimates for our primary model and include selective comparisons using the estimates with the other sets of findings. These added models are intended to be largely checks on the robustness of our main results.

There is an important qualification to our estimates. We do not claim that the HRS sample provides sufficient information to estimate the VSLs for heterogeneous workers over the full range of individual circumstances. As we and others have previously noted (see for example Viscusi and Aldy [2007]), the HRS is intended to reflect behavior of older adults. The younger respondents in the sample are spouses (or partners) of the targeted population. Equally important, the mix of industries and occupations represented may not be as detailed as the samples usually considered for estimating wage risk tradeoffs. Finally, to assure our analysis is aligned with the hedonic wage model's assumptions we limit the sample to respondents paid by the hour. These restrictions yield a smaller sample in comparison to what has been available in other recent hedonic wage studies.

As a result of these considerations, we focus on how the relative risk tradeoffs among workers within this sample vary with health status and longevity threats. Equation (5) adapts the conventional form of a wage hedonic function to include the added terms for our primary model.

$$\begin{aligned}\ln(wage_{it}) &= \beta_0 + \sum_{k=1}^{K}\delta_k \cdot x_{kit} + \beta_1 \cdot BLSrisk_{it} + \sum_{r=1}^{16} a_r \cdot Occ_{rit} \\ &+ \sum_{l=1}^{5}\alpha_l \cdot Age_{lit} \cdot BLSrisk_{it} + \sum_{j=2}^{4}\gamma_j \cdot Tol_{ji(t-1)} \cdot BLSrisk_{it} \\ &+ \theta_1 \cdot Exhealth_{i(t-1)} \cdot BLSrisk_{it} + \theta_2 \cdot Dead_{i(t+1)} \cdot BLSrisk_{it} \\ &+ \theta_3 \cdot Exhealth_{i(t-1)} \cdot Dead_{i(t+1)} \cdot BLSrisk_{it} + \sigma\sum_{s=1}^{6}\rho_s \cdot S_{sit} + \varepsilon_{it}\end{aligned} \tag{5}$$

Individual characteristics including race, gender, and education are represented by x_{kit}, with the subscript t indicating the period of interest. The Occ_{rit} terms identify a set of occupational dummy variables. Our notation uses $(t-1)$ as a subscript to imply the interview wave preceding the occupational outcome (the preceding interview actually occurred two years earlier than the current wave) and $(t+1)$ for the subsequent wave (two years later).

$BLSrisk_{it}$ measures the number of fatalities per 10,000 workers associated with the individual's industry, based on the 2-digit SIC code, and age range (16-19, 20-24, 25-34, 35-44, 45-54, 55-64, 65+).[7] The age- and industry-specific fatality data are from the BLS Census of Fatal Occupational Injuries while the age- and industry-specific employment data are annual averages based on the Current Population Survey.

Age, in our primary model, is treated non-parametrically with the Age_{lit} terms. The five terms classify respondents into one of six age classes, 44 and under, 45-50, 51-55, 56-60, 61-65, and over 65 (i.e., $Age_{lit} = 1$ if respondent *i*'s age falls in the *l*th age class and 0 otherwise; the omitted category is greater than 65). The $Tol_{ji(t-1)}$ terms refer to dummy variables formed using a qualitative variable defined by the risk tolerance measure described in Barsky et al. [1997]. The risk tolerance index is based on responses from a previous interview to questions about choices between a secure job for life and another job with a 50-50 chance of two different income levels.[8] With locally constant relative risk aversion, the answers classify respondents into one of four risk tolerance categories, with category 1 indicating the *least* risk averse and 4 the *most*. The risk

[7] Note that the two youngest age ranges are not represented in our sample.
[8] The risk tolerance question was not asked in all waves of the HRS.

tolerance index included in equation (5) is therefore a measure of financial risk aversion. While by definition an increase in risk aversion defined as in the term's colloquial use increases the VSL, Eeckhoudt and Hammitt [2004] show that the effect of an increase in financial risk aversion on the WTP to reduce mortality risk is, in general, ambiguous.

The last three interaction terms provide our basis for distinguishing the effects of health status/quality of life and serious threats to longevity on the estimated risk tradeoffs. In constructing these terms, we make use of two additional binary variables, represented in equation (5) by $Exhealth_{i(t-1)}$ and $Dead_{i(t+1)}$.

Our quality of life measure, $Exhealth_{i(t-1)}$, is defined based on each individual's response to a question about the state of his or her current health as excellent, very good, good, fair, or poor and exploits the panel featuring the HRS. We use each individual's response to this health question in the preceding wave.[9] Thus the qualitative variable $Exhealth_{i(t-1)}$ corresponds to answers from the interview preceding the observed occupational choice. If a respondent answered that he or she was in excellent health the variable is coded as unity. Otherwise, $Exhealth_{i(t-1)}$ takes a value of zero.

$Dead_{i(t+1)}$ is constructed to identify those individuals facing impending, serious threats to longevity. While information about immediate, serious health conditions may be known to respondents, it appears that they do not report all aspects of their subjective assessments at the time of each wave's interview (see Smith et al. [2001]). Therefore, we use the occurrence of death in the near future to proxy for such information. $Dead_{i(t+1)}$ is a dummy variable to identify people who died between the interview wave when their labor market choices were reported and the subsequent interview. These are confirmed deaths and not sample attrition for other reasons.[10] $Exhealth_{i(t-1)}$ and $Dead_{i(t+1)}$, through the final three interaction terms in equation (5), provide instruments for identifying the effects of excellent quality of life and serious longevity threats on estimated risk tradeoffs.

As noted in Smith et al. [2004], selection into the labor force or into other employment categories such as retirement is an important consideration when examining the occupational choices of older workers. In order to address this, we follow the same logic as discussed in Smith et al. [2004], relying on the Bourguignon et al. [2001] generalization of the two step selection framework. We first estimate a multinomial logit selection model of labor force status. The S_{sit} terms in equation (5) are then the expectations for the wage model's error conditional on employment state that arise from the selection model (see Smith et al. [2004] for further discussion) and reflect choices at the extensive margin (e.g., work, retire, unemployed, etc.) that give rise to selection effects on wage models.

Our wage model estimates are based primarily on information obtained in wave 2 of the Health and Retirement Study (HRS), conducted in 1994. The HRS is a panel survey designed to represent individuals between ages 51 and 61 in 1992 and their spouses (whose ages need not fall within the targeted age range). The initial interviews were conducted in 1992 with follow-up interviews scheduled every two years. The baseline survey included 12,652 persons (7,600 households). The panel was designed to

[9] Both current and previous wave responses to these health condition questions are also significantly different for respondents who die between waves 2 and 3 and those who do not (results for the current wave are: χ^2=40.40, p-value=0.00 and for the previous wave: χ^2=21.73, p-value=0.00).

achieve several objectives (see Juster and Suzman [1995]) including monitoring work, health, and income relationships. The HRS structure allows construction of the quality of life instrument based on wave 1 (1992) subjective health status and the impending serious health shock instrument based on deaths between waves 2 (1994) and 3 (1996). For consistency with the notation adopted in equation (5), the time subscripts $t-1$, t, and $t+1$ correspond to 1992, 1994, and 1996 respectively.

While a number of waves are available for analysis, when considering waves beyond wave 2 (1994), the analysis confronts the limitations imposed by the age profile of the sample. This structure implies a decrease in the number of respondents with typical employment patterns. This decline limited our earlier analysis of wage/job risk tradeoffs as the HRS sample aged.[11] For this reason, our current analysis focuses on labor market experience as of wave 2. As noted earlier, our sample for the hedonic wage model is limited to workers paid on an hourly basis, where the BLS fatality rates are more likely to reflect the job risks faced by workers.

Estimation of the multinomial logit selection model requires that we define relevant labor force status categories, working and paid by the hour being one such category. We select five other categories: working and not paid by the hour, unemployed, retired and not currently working, disabled, and not in the labor force. As in Smith et al. [2004], we hypothesize that an individual's marital status, age, number of people living in his household, his household's capital income, and whether or not his health limits his work opportunities will influence his labor force status decision.

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[10] Unfortunately the records do not include information about the causes of death.

[11] See Smith et al. [2004] for an example of the marked decline in workers from wave 1 continuing to work by wave 4.

Table 1 reports the results of the multi-state selection model in addition to summary statistics by labor force status. Coefficients are interpreted as relative to the excluded category, working and not paid hourly. For example, the coefficient on age is positive and significant for the disabled category, suggesting that, relative to respondents who are working and not paid hourly, respondents who are disabled tend to be older. The results suggest that relative to respondents who are working and not paid hourly, respondents who are working and paid hourly, unemployed, or disabled are less likely to be married. The negative and significant coefficient on household capital income for all categories indicates that household capital income is lower for other respondents relative to respondents who are working and not paid hourly.[12] Relative to respondents who are working and not paid hourly, respondents who are unemployed, retired and not working, disabled, or not in the labor force are more likely to report that their health limits their ability to work. As expected, this variable is insignificant for the working and paid hourly category. In general, the results of the multi-state selection model are consistent with expectations. Based on these results, we estimate conditional expectations for the wage equation's error for each labor force state and include these terms as independent variables in the hedonic wage model.

Table 2 provides summary statistics and a selected set of parametric estimates from six of the hedonic wage models motivated by equation (5).[13] All estimated wage

[12] Household capital income is the sum of household business or farm income, self-employment earnings, business income, gross rent, dividend and interest income, trust funds or royalties, and other asset income.
[13] In order to assure the existence of a wage-risk premium in line with the existing literature, we estimate a conventional hedonic wage model (equation (5) without any of the interaction terms). The estimated coefficient on the risk term in this equation is statistically significant and equal to 0.042. This implies an estimated VSL of $9.6 million (2000$), above the median of $7 million reported in Viscusi and Aldy's [2003] meta analysis, but still within the range of values from what they considered to be the most reliable studies.

equations were adjusted for selection effects as discussed above. While not reported, the selection terms are jointly significant influences on wages (p-value = 0.01). The number in parentheses below the estimated parameters in the second through seventh columns of Table 2 presents the ratios of the estimated coefficients to their respective estimated standard errors based on the Huber [1967] robust covariance matrix. In earlier research we compared the Huber robust covariance matrix along with bootstrapped standard errors and found no differences in conclusions about the role of job risk, age, and risk preferences. Each of the models includes fixed effects for the occupational categories identified in equation (5) but these are not reported in the table.

The remaining models in the table are distinguished based on the treatment of the age terms and the inclusion or exclusion of the quality of life and subjective longevity interaction terms. Specification (2) includes age in linear and quadratic terms as well as in categorized variables that are interacted with the BLS risk. Specifications (3) and (4) drop the categorized treatment of age and use the linear and quadratic terms with risk interactions as in work by Aldy and Viscusi. These two specifications vary based on the inclusion or exclusion of risk tolerance, quality of life and longevity threat effects. In order to investigate whether the treatment of serious longevity threats influences estimates of risk tradeoffs independent of quality of life considerations, we re-estimate the primary model, eliminating the interaction effects used to identify the quality of life assessments. We repeat the same exercise eliminating interactions for longevity threats to consider whether the quality of life effects are robust to this exclusion. The final two specifications in Table (2) report these results. In addition to the estimates reported in Table 2, we also considered a variety of other models that varied the treatment of the risk tolerance, the quality of life/health status, and the longevity effects terms. All of these variations were developed to evaluate whether our basic conclusions would change with plausible variations in the model specification. A summary of these comparisons would suggest that our findings are quite robust to all of the model variations we considered.

Because the BLS risk term enters alone and in interactions with age, risk tolerance measures, and the variables used to gauge the quality of life and longevity perceptions, the size, magnitude and significance of the coefficient for the simple risk term cannot be interpreted independently. The interaction terms including the risk measure must be considered in evaluating the size and significance of the estimated marginal effect of job risk on the natural log of wages. To test for the effects of risk on wage compensation requirements we need to consider several different coefficient sums. The estimated wage differentials are formed by summing the coefficients on the risk term and the appropriate interaction terms for each group. For example, in order to test the hypothesis that $\frac{\partial(\ln(wage))}{\partial(BLSrisk)}$ is zero for respondents in the second age and risk tolerance categories, who reported poor health two years earlier, and who die in the next two years, the relevant test statistic is given by $\hat{\beta}_1 + \hat{\alpha}_2 + \hat{\gamma}_2 + \hat{\theta}_2$.

In discussing the results of our hedonic wage analysis, we begin with the primary model. The first step in considering the effects of age, risk tolerance, quality of life/health status, and anticipated longevity on estimated wage-risk tradeoffs is to test whether all these interactions can jointly be omitted from the model. We reject the null hypothesis that the interaction terms are jointly zero (p-value = 0.006). The set of age and risk tolerance interactions terms are jointly significant at a marginal level (p-value =

0.099); the health status and anticipated longevity effects strengthen the importance of the set of interactions.[14]

To highlight the effects of quality of life and serious longevity threats on the estimated risk tradeoffs, Table 3 presents a subset of our results from four separate models. In particular, we select the two most risk averse categories and the three age categories that span the age range of which the HRS is intended to be representative, 51-55, 56-60, and 61-65. Each cell in Table 3 reports the estimated coefficient sum, $\frac{\partial(\ln(wage))}{\partial(BLSrisk)}$ by group. In parentheses below each estimated coefficient sum, we report the p-value for a test that the sum is significantly different from zero. In columns 2 through 5 we report the coefficient sums for two models, each of which includes terms for both subjective health-related quality of life, under "Excellent health" and "Not excellent health", and the presence and absence of threats to longevity, under "Longevity threat" and "No longevity threat".[15] The coefficient sum reported in the left-hand side of each cell results from specification (1), our primary model. The coefficient sum reported in the right-hand side of each cell results from specification (3). Recall these specifications differ only in their treatment of age; specification (1) treats age non-parametrically while specification (3) follows the parametric treatment of age used in

[14] While not reported, we estimated a model excluding the age and risk tolerance interaction terms. The estimated coefficients (and levels of significance) on the three interaction terms measuring the effects of health status and longevity threats are similar to our primary model.

[15] Relative to those results presented here, we were unable to reject the hypothesis that $\frac{\partial(\ln(wage))}{\partial(BLSrisk)}$ is zero for some of the groups not represented in Table 3. However, all unreported coefficient sums from this model are either positive and statistically significant or not statistically different from zero. See Smith et al. [2004] for a discussion of differences between the younger workers represented in the HRS sample and the Current Population Survey. Full results are available from the authors.

Viscusi and Aldy.[16] Computation of comparable coefficient sums using the results of specification (3) required that we focus on particular ages, rather than age ranges. For comparison, we computed coefficient sums for ages 55, 60, and 65. Comparing the terms within each cell, we see the magnitude of the estimated effects are very similar regardless of our treatment of age. Given these similarities, we restrict our attention in the discussion that follows to the results for specification (1), reported in the left-hand side of each cell.

We conduct pair-wise comparisons of the estimated coefficient sums (from specification (1)), testing whether or not the observed differences across groups are statistically significant. The null hypothesis is equality of the estimated coefficient sums corresponding to each comparison pair. Table 4 reports the p-values of a subset of these tests where the age and risk tolerance categories are held constant.

The results allow us to distinguish three types of effects. *First*, we discuss test results for how impending serious longevity threats -- *facing the prospect of death* -- influence the risk tradeoff. This relationship also serves as a test of one interpretation of the Pratt-Zeckhauser [1996] dead anyway hypothesis—individuals facing impending serious health threats would be willing to pay more to realize even a small risk reduction. *Second*, we investigate whether an individual's subjective opinion of his health-related

[16] The results of specification (3) are consistent with an inverted U-shaped relationship between age and the VSL as has been found in work summarized by Aldy and Viscusi [2007b]. However, the estimated peak of the relationship occurs around age 56 for our sample. Aldy and Viscusi report a peak at about age 40. Two aspects of this comparison should be noted. First, if we review the results from our earlier analysis that did not take account of quality or quantity issues and introduced age non-parametrically, the estimated wage rsik tradeoffs were not precisely estimated for the 61-to 65 age group. Age category 56-60 was the last group with significant estimates of the effects by risk tolerance group (see table 3 in Smith et al. [2004]). Second and equally important, a quadratic specification for any variable restricts effects to a parabolic form (either U or inverted U shapes). A non-parametric or categorical coding allows for different effects by age category but does not restrict the form of those differences. As a result, it is our preferred treatment for examining age effects.

quality of life influences the relative size of the effect due to serious longevity threats. *Finally*, we examine the independent effect of quality of life on the estimated risk tradeoffs.

Pair wise comparisons of the terms in the left-hand sides of columns 2 and 3, and 4 and 5 respectively provide evidence consistent with an effect of longevity threats on risk tradeoffs and support what we described as the dead anyway hypothesis. For a given age, risk tolerance category, and quality of life, the dead anyway hypothesis would suggest larger coefficient sums among individuals who expect serious adverse health shocks (columns 2 and 4 of Table 3). For people who rated their health as excellent in wave 1 with an impending serious longevity threat, the coefficient sums presented in column 2 are at least six times the size of the estimated risk tradeoff for those in excellent health without the impending threat. This difference is significant at the high end of conventional levels (p-value = 0.085). Comparisons of columns 4 and 5 confirm the large differential associated with impending threats to longevity for individuals in health categories other than excellent. This difference is also significant (p-value = 0.050). The estimated coefficient sums reported in columns 2 and 4 are large and would imply extremely large VSL measures. We return to this point later in the discussion of our results.

Comparisons of coefficient sums associated with excellent or not excellent health within the same age, risk tolerance, and longevity threat category allow us to examine quality of life effects. Among those respondents with impending serious longevity threats, the coefficient sums for the excellent health categories (in column 2) consistently exceed those for the “not in excellent health” groups (in column 4). However, as reported in Table 4, these differences are not statistically significant (p-value = 0.254). Comparing the coefficient sums in columns 3 and 5 yields a smaller but statistically significant quality of life premium among respondents who do not expect serious health threats (p-value = 0.018).

In order to investigate whether the effects of serious longevity threats displayed a fairly robust influence on risk tradeoffs independent of quality of life considerations, we conduct a similar exercise using the results of specification (5). A subset of these results is presented in columns six and seven of Table 3. Pair-wise comparisons confirm a significant premium (p-value = 0.019) for those who face serious threats, the relative size of which is consistent with the model accounting for quality of life differences. However, the coefficient sums are not as large, suggesting that quality of life and serious threats to longevity may well be confounded in these estimates.

The results presented in the final two columns of Table 3 reflect the impact of differences in quality of life, independent of longevity threats, on the estimated wage-risk tradeoffs. The selected coefficient sums are formed using the results of specification (6), which excludes the interaction terms used to identify the effects of serious longevity threats. The coefficient sums in the eighth column, reflecting those with excellent health-related quality of life, are approximately twice as large as the respective coefficient sums in the final column. The estimated quality of life premium suggested by pair-wise comparisons of these coefficient sums is significant (p-value = 0.018).

To further explore the implications of allowing the estimated risk tradeoffs to vary with quality of life, we compute what we call relative risk tradeoffs for the age/risk tolerance groups with significant coefficient sums in the final two columns of Table 3.

Table 5 presents these relative risk tradeoffs, which represent ratios of the associated VSL estimates in 2000 dollars for respondents in excellent health to the VSL estimates for respondents not in excellent health. Below each relative tradeoff is the p-value for a test that the ratio is significantly different from one. We test this hypothesis using the nlcom command in Stata.[17] In each case, the relative tradeoff is significantly different from one at a p-value of 0.05 or less, indicating that individuation on the basis of quality of life leads to a significant difference in estimates of the marginal rate of substitution between risk and wealth. For the respondents in our sample, excellent health almost doubles the incremental risk tradeoff.

As we noted, if differences in health status lead to variation in the degree of complementarity between consumption and labor supply, we should also expect differences in estimated risk tradeoffs with health status. Our results are consistent with this response. Of course, we would not suggest that all the variation in VSL estimates with age can be attributed exclusively to differences in the health status of older adults. Rather we argue that it is important to acknowledge that the older adults in our sample differentially experience serious health conditions that can alter their subjective assessments of the quality of their lives and influence their labor/leisure choices.

For each individual, consumption decisions are constrained by time as well as income; time is required to enjoy the consumption that is permitted by increased income arising from the labor each person supplies. Chetty's analysis reminds analysts that when we attempt to understand the interactions between labor supply and risk preferences revealed through wage/risk tradeoffs, we have several "windows" on preferences. In our

[17] The calculation assumes that wage, hours worked per week, and weeks worked per year are fixed at the mean values for the respective age and risk tolerance category.

analysis variation in health status was one means to reconcile Chetty's argument for different degrees of complementarity between labor supply and consumption to motivate variation in the elasticity of the marginal utility of consumption (and in the labor supply elasticity and VSL).

Before turning to a brief discussion of the implications of these findings, it is important to acknowledge that our results are consistent across age and risk tolerance classes and different estimated wage models. We have limited the discussion of the full structure of our wage model or the multiple state selection model because both have been discussed for another closely related sample in Smith et al. [2004]. However, these other components of the model assure that for this sample the effects of the quality of life and serious longevity threat terms are being estimated within a framework that accounts for an array of conventional covariates as well as the selection effects associated with the labor force participation decision.

As a final exercise, we estimate two additional models, both variants of our primary model, in order to examine the robustness of our results to changes in our longevity threat measure. First, we replace the future deaths variable with an alternative instrument measuring serious threats to longevity, a qualitative variable identifying a newly diagnosed cancer or stroke between waves 2 and 3. Compared to the number of deaths experienced between waves 2 and 3, there were more new cases of these two serious health conditions (i.e., 46 cases versus 23 deaths).[18] We re-estimated the model reported in columns six and seven of Table 3 using this new variable. As a second exercise, we replace the future deaths variable with a variable indicating more than one hospital stay between waves 2 and 3.[19] Table 6 reports the resulting coefficient sums from these two models for a subset of age and risk tolerance groups.[20] For both models, pair-wise comparisons indicate significant differences between the coefficient sums reported in columns 2 and 3 (p-value = 0.078), and 4 and 5 (p-value = 0.018). The results confirm a premium when we replace a future death with these two alternative indicators of serious threats to longevity.

V. Implications

The mature workers in our sample appear to select jobs with risks in ways that are conditioned by their personal assessments of their longevity and the quality of their lives. Serious threats to longevity affect the expected quantity of life remaining and, as a result, the relative risk tradeoffs. We have provided a model that indicates a clear role for subjective assessments of health status in these tradeoffs as well. Those in excellent health require significantly greater compensation to accept risk increases compared to those who are not. This finding lends support to health status or quality of life as a potential source of variation in the degree complementarity between consumption and labor supply.

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[18] The number of deaths may seem especially limiting, considering all the coefficients reported in Table 1. However, only two parameters (θ_2 and θ_3) are estimated to distinguish the hypotheses regarding quality of life and longevity threats.

[19] 124 respondents report more than one hospital stay between waves 2 and 3. In these final two models, the sample sizes in the hedonic wage model reduce to 2521 and 2520 due to the exclusion of those who die between waves 2 and 3 and to those respondents for whom we are unable to determine whether or not they have experienced a stroke or cancer or a hospital stay in between waves 2 and 3. While not reported, the estimated coefficients of the selection models and the hedonic wage models are similar to the results reported in Tables 1 and 2.

[20] All but one unreported coefficient sum for the final two models is either positive and statistically significant or not statistically different from zero. In the future cancer or stroke model, the estimated coefficient sum for the older than 66, most risk tolerant with no future cancer or stoke group is negative and statistically significant.

Thus, these estimates suggest that current modeling decisions that rely on simple descriptions of each individual's life expectancy based on age alone may offer incomplete basis for judging the wage/ risk tradeoffs underlying the measures for the VSL. We must consider whether a person's perceived health status influences the mix of consumption goods and services selected, the amount and quality of leisure time selected, and the net outcome of these factors on the measured extent of complementarity between the consumption and labor supplied to the market. These influences may become more relevant as a person ages.

Our findings support the feasibility of using a sample of individuals in a panel format to sort out some of the determinants of heterogeneous risk preferences, rather than resorting to experts' judgments as in the QALY literature. The detailed records of actual health conditions, subjective evaluations of perceived health status, and panel structure of the HRS allow us to demonstrate how it would be possible to measure the influence of these factors on the relative size of different individuals' risk tradeoffs. The next step is to expand the samples considered beyond the older respondents in the HRS and evaluate where analogous serious longevity threats and quality of life judgments affect younger adults in similar ways.

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Table 1. Multinomial logit selection model of labor force status[a]

Variable	**Working and paid hourly**		**Unemployed**		**Fully retired or partially retired but not working**		**Disabled**		**Not in labor force**	
	Coefficient (z-statistic)	Mean (standard deviation)	Coefficient (z-statistic)	Mean (standard deviation)	Coefficient (z-statistic)	Mean (standard deviation)	Coefficient (z-statistic)	Mean (standard deviation)	Coefficient (z-statistic)	Mean (standard deviation)
Married	-0.30* (-4.51)	0.77 (0.42)	-0.97* (-6.68)	0.64 (0.48)	-0.041 (-0.53)	0.78 (0.41)	-0.75* (-5.99)	0.57 (0.50)	0.22* (2.04)	0.86 (0.34)
Age	-0.0042 (-0.86)	55.91 (5.03)	-0.024* (-2.03)	55.38 (5.86)	0.18* (28.89)	60.49 (5.25)	0.037* (3.05)	56.94 (4.83)	-0.014* (-2.10)	55.42 (5.55)
# people in household	0.046* (2.12)	2.62 (1.30)	0.15* (3.36)	2.72 (1.41)	-0.055* (-2.05)	2.39 (1.14)	0.75 (1.76)	2.51 (1.46)	0.15* (5.42)	2.86 (1.42)
Household capital income	-0.000033* (-14.32)	3301.05 (12359.68)	-0.000015* (-3.42)	4801.75 (23411.40)	-0.000010* (-7.75)	5971.67 (26052.8)	-0.000075* (-5.13)	932.50 (7595.66)	-0.0000021* (-2.32)	13239.34 (80454.62)
Health limits work	-0.016 (-0.19)	0.10 (0.30)	1.37* (9.43)	0.32 (0.47)	2.29* (31.62)	0.51 (0.50)	5.63* (18.84)	0.97 (0.16)	1.35* (14.98)	0.28 (0.45)
Constant	0.23 (0.79)		-1.13 (-1.60)		-11.45* (-29.49)		-7.04* (-9.24)		-1.36* (-3.40)	
N	2708		261		2635		436		1013	

[a] * indicates statistical significance at the 5% level. Reference category, with 3879 observations, is working and not paid hourly. Summary statistics for this category are as follows: Married—0.83 (0.37), Age—56.04 (5.34), # people in household—2.58 (1.14), Household capital income—20841.68 (80357.23), Health limits work—0.09 (0.29). Total number of observations for the multinomial logit selection model is 10932 with a pseudo R-squared value of 0.15.

Table 2. Hedonic wage model: Selected parameters[a]

Independent variable	(1)	(2)	(3)	(4)	(5)	(6)	Mean (standard deviation) or fraction of sample
Male	0.268* (12.40)	0.2698* (12.43)	0.2667* (12.42)	0.2682* (12.46)	0.2689* (12.43)	0.2680* (12.39)	0.42
White	-0.0158 (-0.80)	-0.0168 (-0.85)	-0.0146 (-0.74)	-0.0130 (-0.66)	-0.0141 (-0.71)	-0.0173 (-0.88)	0.78
Years of education	0.0268* (8.16)	0.0266* (8.08)	0.0267* (8.14)	0.0268* (8.21)	0.0273* (8.33)	0.0264* (8.06)	11.58 (2.72)
BLS risk	-0.0641 (-1.56)	-0.0451 (-1.06)	-2.4423* (-3.09)	-2.431* (-3.01)	-0.0602 (-1.47)	-0.0657 (-1.60)	0.64 (0.63)
Age	–	0.0449 (-1.72)	–	–	–	–	55.91 (5.03)
Age^2	–	-0.0004 (-1.71)	–	–	–	–	–
Age*BLS risk	–	–	0.0856* (3.15)	0.0882* (3.25)	–	–	–
Age^2*BLS risk	–	–	-0.0007* (-3.19)	-0.0008* (-3.34)	–	–	–
Age_1*BLS risk	-0.0810 (-0.67)	-0.0115 (-0.08)	–	–	-0.0739 (-0.61)	-0.0784 (-0.65)	0.027
Age_2*BLS risk	-0.00444 (-0.06)	-0.0213 (-0.30)	–	–	0.0003 (0.004)	-0.0031 (-0.04)	0.081
Age_3*BLS risk	0.0532 (1.20)	0.0271 (0.58)	–	–	0.0597 (1.35)	0.0562 (1.27)	0.35
Age_4*BLS risk	0.0733 (1.87)	0.0510 (1.23)	–	–	0.0782* (1.99)	0.0724 (1.84)	0.37
Age_5*BLS risk	0.0679 (1.65)	0.0542 (1.29)	–	–	0.0727 (1.77)	0.0673 (1.64)	0.15
$Tol_{2(t-1)}$*BLS risk	0.00815 (0.23)	0.0077 (0.22)	0.0069 (0.20)	–	0.0113 (0.32)	0.0086 (0.24)	0.11
$Tol_{3(t-1)}$*BLS risk	0.0318 (0.93)	0.0337 (0.98)	0.0288 (0.84)	–	0.0389 (1.14)	0.0383 (1.12)	0.11
$Tol_{4(t-1)}$*BLS risk	0.0524* (2.11)	0.0532* (2.15)	0.0500* (2.01)	–	0.0531* (2.14)	0.0541* (2.18)	0.66
$Exhealth_{(t-1)}$*BLS risk	0.0514* (2.36)	0.0522* (2.39)	0.0513* (2.36)	–	–	0.0515* (2.37)	0.23
$Dead_{(t+1)}$*BLS risk	0.226* (1.96)	0.2238 (1.94)	0.2240 (1.94)	–	0.2575* (2.35)	–	0.085

[a] The numbers in parentheses in the second column are the ratios of the estimated coefficients to their estimated standard errors based on the Huber [1967] robust estimates for the variance covariance matrix. * indicates statistical significance at the 5% level. The final column reports means and standard deviations for the continuous variables, or the fraction of the sample for the dichotomous variables listed in column one. The values reported in this column for the interaction terms represent the fraction of the sample in the relevant age or risk tolerance category (independent of the job risk term). The values reported for $Exhealth_{(t-1)}$ and $Dead_{(t+1)}$ designate the fraction of the sample reporting excellent health in the preceding wave and the fraction of the sample who died between wave 2 and wave 3.

Exhealth$_{(t-1)}$*Dead$_{(t+1)}$ *BLS risk	0.3527 (0.99)	0.3498 (0.99)	0.3484 (0.98)	–	–	–	–
Number of observations	2708	2708	2708	2708	2708	2708	2708
R^2	0.40	.372	.372	.368	0.370	0.370	–

Table 3. Selected tests of quality of life and longevity threat effects on estimated risk tradeoffs[a]

Risk tolerance category	Risk tradeoff with quality of life and longevity threat adjustments				Risk tradeoff with longevity threat adjustment		Risk tradeoff with quality of life adjustment	
	Excellent health		Not excellent health		Longevity threat	No longevity threat	Excellent health	Not excellent health
	Longevity threat	No longevity threat	Longevity threat	No longevity threat				
Age 51-55								
3	0.651* | 0.660* (0.052) | (0.049)	0.072* | 0.088* (0.052) | (0.009)	0.247* | 0.260* (0.037) | (0.027)	0.021 | 0.036 (0.544) | (0.236)	0.296* (0.008)	0.038 (0.256)	0.080* (0.030)	0.029 (0.404)
4	0.672* | 0.681* (0.046) | (0.043)	0.093* | 0.109* (0.002) | (0.000)	0.268* | 0.282* (0.022) | (0.015)	0.042* | 0.058* (0.095) | (0.003)	0.310* (0.005)	0.053* (0.031)	0.096* (0.001)	0.046* (0.073)
Age 56-60								
3	0.671* | 0.659* (0.046) | (0.049)	0.092* | 0.087* (0.006) | (0.011)	0.288* | 0.259* (0.014) | (0.029)	0.062* | 0.035* (0.002) | (0.245)	0.315* (0.005)	0.057* (0.058)	0.096* (0.004)	0.045 (0.143)
4	0.692* | 0.680* (0.040) | (0.043)	0.113* | 0.108* (0.000) | (0.000)	0.262* | 0.280 (0.030) | (0.016)	0.036 | 0.056* (0.350) | (0.002)	0.329* (0.003)	0.071* (0.000)	0.112* (0.000)	0.061* (0.002)
Age 61-65								
3	0.666* | 0.621* (0.048) | (0.065)	0.087* | 0.048 (0.041) | (0.230)	0.247* | 0.221* (0.037) | (0.067)	0.021 | -0.0032 (0.544) | (0.930)	0.309* (0.007)	0.051 (0.171)	0.091* (0.025)	0.040 (0.295)
4	0.687* | 0.642* (0.042) | (0.057)	0.108* | 0.069* (0.001) | (0.036)	0.268* | 0.242* (0.022) | (0.042)	0.042* | 0.018 (0.095) | (0.507)	0.323* (0.004)	0.066* (0.021)	0.107* (0.001)	0.056* (0.053)

[a] The numbers are coefficient sums defining $\partial(\ln(wage))/\partial(BLSrisk)$ for each age, risk tolerance, quality of life, and severe health threat groups. Numbers in parentheses are p-values for the null hypothesis that the coefficient sum is zero. * indicates p-value of 0.10 or lower. The Huber [1967] robust estimate for the variance covariance matrix of the estimated parameters is used in forming the variance of the coefficient sums implied by each element in the table.

Table 4. Results of pair-wise comparisons[a]

	Excellent, No longevity threat	Not excellent, Longevity threat	Not excellent, No longevity threat
Excellent, Longevity threat	0.085	0.254	0.061
Excellent, No longevity threat		0.134	0.018
Not excellent, Longevity threat			0.050

[a] The table reports p-values for pair-wise comparisons of the estimated wage-risk tradeoff, $\partial(\ln(wage))/\partial(BLSrisk)$, for each quality of life and severe longevity threat group for a given age and risk tolerance category based on the results presented in Tables 1, 2, and 3. In all cases the Huber [1967] robust estimate for the variance covariance matrix of the estimated parameters is used in forming the test statistic. The null hypothesis is equality of the estimated coefficient sums for each pair presented in the table.

Table 5. Relative risk tradeoffs for quality of life effects[a]

Age category	Risk tolerance category	Relative risk tradeoff (p-value)
51-55	4	2.16* (0.014)
56-60	4	1.85* (0.000)
61-65	4	1.93* (0.004)

[a] The relative risk tradeoffs are formed using estimates of the average wage, hours, and weeks worked for each age risk tolerance group. Numbers in parentheses are p-values for the null hypothesis that the relative risk tradeoff is one. * indicates that the ratio is significantly different from one with a p-value of 0.05 or less using the nlcom command in Stata. The test assumes wage, hours worked per week, and weeks worked per year are fixed at their mean values for the respective age and risk tolerance category.

Table 6. Tests of effects of longevity threat hypothesis with alternative measures of baseline mortality risk[a]

	Baseline mortality measure: newly diagnosed future cancer or stroke		Baseline mortality measure: more than one future hospital stay	
Risk tolerance category	Future cancer or stroke	No future cancer or stroke	More than one future hospital stay	Zero or one future hospital stay
Age 51-55				
3	0.116* (0.044)	0.035* (0.316)	0.134* (0.013)	0.033* (0.348)
4	0.142* (0.006)	0.061* (0.015)	0.162* (0.001)	0.061* (0.017)
Age 56-60				
3	0.137* (0.012)	0.056 (0.073)	0.154* (0.003)	0.053 (0.092)
4	0.163* (0.001)	0.082* (0.000)	0.182* (0.001)	0.081* (0.000)
Age 61-65				
3	0.130* (0.026)	0.049 (0.214)	0.144* (0.008)	0.042 (0.283)
4	0.156* (0.003)	0.075* (0.013)	0.172* (0.000)	0.071* (0.021)

[a] The numbers are coefficient sums defining $\partial(\ln(wage))/\partial(BLSrisk)$ for each age, risk tolerance, and baseline risk groups. Numbers in parentheses are p-values for the null hypothesis that the coefficient sum was zero. * indicates statistical significance at the 5% level. The Huber [1967] robust estimate for the variance covariance matrix of the estimated parameters is used in forming the variance of the coefficient sums implied by each element in the table.