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Attitudes towards large income risk in welfare states: an international comparison

BY

Fred Schroyen AND **Karl Ove Aarbu**

Norges Handelshøyskole

NORWEGIAN SCHOOL OF ECONOMICS

Attitudes towards large income risk in welfare states: an international comparison*

Fred Schroyen† and Karl Ove Aarbu‡

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Abstract Using survey data and the instrument developed by Barsky *et al.* (1997), we estimate the distribution of attitudes towards income risk in a country where many employment and health-related risks are generously covered by a tax financed social insurance system (Norway 2006) . Under a CRRA assumption, the sample average for the coefficient of relative risk aversion is 3.8 with a standard deviation of 2.3. This number is then contrasted to that for five other OECD countries where risk attitudes have been measured using the same instrument and also prior to the financial crisis: Chile, France, Italy, The Netherlands and the US. When we relate this distribution for stated relative risk aversion to that for generosity of social insurance and the risks related to employment and health expenditure, a picture emerges suggesting that more extensive welfare states induce higher risk tolerance for foreground risks–a relationship that is in line with the theory on risk vulnerability.

JEL classification D12; D81

Keywords risk aversion; stated preferences; income lotteries; background risk; risk vulnerability; welfare state.

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†Dept of Economics, NHH Norwegian School of Economics, N-5045 Bergen; e-mail: `fred.schroyen@nhh.no`.

‡Tryg Forsikring and Dept of Economics, NHH Norwegian School of Economics, N-5045 Bergen; e-mail: `karl.ove.aarbu@tryg.no` or `karl.aarbu@nhh.no`.

1 Introduction

Many big decisions in life are made behind a veil of uncertainty. These include savings decisions, career choices, investments in human and health capital, choices of lifestyle—even the most informed decisions yield payoffs that are contingent on states of the world of which the realisation is uncertain. Among the reasons why individuals make some decisions rather than others, is their attitude toward risk, formalized by Pratt (1964) and Arrow (1965) through the concepts of absolute and relative risk aversion. Many of these decisions are also made with independent risks looming in the background. It is one of the virtues of a welfare state to mitigate some of these background risks through social insurance. An unemployment insurance with a high replacement rate financed through proportional income taxation will cushion an employee against unemployment risk and make her more willing to take on a job in a sector exposed to the business cycle. A citizen enrolled in a tax financed health care system with low co-payments need not to worry about medical care bills taking a large chunk out of disposable income in case of ill-health, neither about insurmountable health insurance premia in case of a job loss or in case illness of oneself or a close relative forces one out of the labour market. In this paper, we are interested in citizens' attitudes towards *large* income risks and how welfare state arrangements can moderate the aversion towards such risks by providing a safety net against other uninsurable risks related to human and health capital and other non-marketable assets against which it is hard to find private insurance.

In contrast to many studies that inquire about attitudes towards small risks, elicitation of preferences w.r.t. large risks means that one is methodologically restricted to asking people about their preferences over *hypothetical* lotteries. Since we would like the instrument to be easily comparable across countries and over time, it is also an advantage that the measure is unit-free. An instrument that meets these criteria is the set of hypothetical life time lotteries introduced by Barsky, Juster, Kimball and Shapiro (1997)–BJKS hereafter. Moreover, since we would like to verify whether such preferences may be affected by different welfare state arrangements, it is desirable to have responses from a representative sample of citizens rather than from a narrow subset (e.g., students). To the best of our knowledge, the BJKS instrument has been used on large samples that are representative for the entire adult population–or a significant part of it–in five OECD countries: Chile (2002), France (2004), Italy (2007), The Netherlands (2004) and the United States (late 1990s and 2002).[1]

[1] In France and the US, the BJKS questions are at regular time intervals included in nationwide surveys. The instrument has also been applied on a sample of 2619 Croatian retirees (Brown *et al.* 2015) in the period November 2008-January 2009. However, we disregard this 'observation' for two reasons. First, the survey was carried out after the onset of the financial crisis (in

In this paper, we first complement this list with a sixth OECD country, Norway. In 2006 the BJKS instrument was used in a survey to which a representative sample of around 1500 adult Norwegians responded. Like for the other samples, the elicited measure of risk aversion exhibits a considerable degree of heterogeneity among citizens. Although gender and age explain some of this variation, a considerable portion is still left unexplained. Using our estimated model, and conditional on each respondent's observed answers to the BJKS lottery questions, we impute to him/her a dimensionless cardinal measure of stated risk aversion which is subsequently used as a co-variate in the analysis of different lifestyle and labour market choices.[2] This measure correlates significantly with risk behaviours for which true risk preferences are expected to matter (likelihood for being obese, a regular smoker, employed in the private sector or as a top manager), which we take as a strong indication that it is a valid measure for these peferences.

In the second and main part of the paper, we make the auxiliary assumption of CRRA preferences and calculate the sample average for the stated coefficient of relative risk aversion for Norway and each of the five aforementioned countries. This–admittedly small–sample of average relative risk aversion coefficients ranges from 3.8 (Norway) to 10.2 (Chile). We posit that the risk preferences measured through the BJKS lotteries are conditional on the scope and generosity of the social insurance system in each country, as well as the sizes of background risks which such systems insure against. More specifically, we hypothesise that (i) smaller background risks and (ii) more complete social insurance induce a lower aversion towards 'foreground' income risks. The theoretical mechanism that motivates this hypothesis is that of vulnerability to undesirable background risks.

To give support to our hypothesis, we collect and construct several measures of welfare state generosity and of the risks that welfare state programmes protect against. We then rank the countries in terms of uncovered background risk. From this ranking we infer that the European welfare states leave citizens with lower background risk then those of the US and Chile, with Norwegian citizens possibly facing the lowest of such risks.

The paper unfolds as follows. In the next section, we present and motivate the hypothetical income gamble questions and describe theoretically how we recover our measure of relative risk aversion from the answers given to the hypothetical questions. We also discuss the pros and cons of the elicitation instrument *vs* other

contrast to the surveys for the aforementioned countries); and second, because Croatia not being an OECD member country made it very hard to find comparable figures for the size of its welfare state.

[2] In this respect, our approach is different from BJKS and Kimball *et al.* (2008) who estimate an unconditional cardinal measure of relative risk tolerance, but closer to Sahm (2012), who estimates a conditional measure of relative risk tolerance measure that she subsequently uses to explain risky holdings of assets.

instruments used in the literature. Section 3 presents the Norwegian 2006 survey and the estimates of the unconditional and conditional distribution of stated risk preferences. In section 4, we compare the Norwegian sample mean for relative risk aversion with that for five other OECD countries and relate the differences to the size of the respective social insurance programmes. We decompose these programme sizes in terms of generosity and risk exposures, and propose a country ranking of background risks. Section 5 concludes.

2 The risk aversion elicitation instrument

2.1 Revealed *vs* stated preferences

There are several ways to elicit risk preferences. The traditional and indirect way is to analyze market behaviour, such as portfolio choice (Blume and Friend, 1975, Bucciol and Miniaci, 2011), or insurance contract choice (Drèze, 1981, Szpiro, 1986, Cohen and Einav, 2007); also the size of consumption responses to labour market shocks, combined with labour supply elasticities, have been used to measure risk aversion (Chetty, 2006, de Linde Leonard, 2012). The more recent and direct way is to study non-market behaviour, for example the choices individuals make among risky alternatives in an experiment or when answering a questionnaire. These may be choices between lottery tickets (Barsky *et al.*, 1997, Holt and Laury, 2002, 2005, Kimball *et al.*, 2008, Choi *et al.*, 2007, Harrison *et al.*, 2007, Vieider *et al.* 2015, Falk *et al.* 2017), the willingness to pay for acquiring such tickets if prizes are in the gain domain (Hartog *et al.*, 2002, Guiso and Paiella, 2008, Rieger *et al.* 2015)), the willingness to avoid such tickets if prizes are in the loss domain (Rieger *et al.*, 2015), or the minimum selling price if endowed with such tickets (Becker *et al.* 1964, Kachelmeier and Shehata, 1992). Depending on whether the lotteries are played (and the outcomes paid out) or not, the preferences thus elicited are said to be revealed *vs* stated.[3]

A central issue in the second branch of literature is whether the stated preferences measure something else than revealed preferences. This question has been studied in detail in the experimental literature, especially with the help of the multiple price list instrument (MPL) popularised by Holt and Laury (2002).[4] In

[3] Some studies ask subjects about their attitudes towards risk in a way that has no direct counterpart in terms of a willingness to pay. E.g., in the German Socio-Economic Panel, 22 000 individuals are asked to rate on a 0-10 scale their answer to the question: "How willing are you to take risks, in general?". Dohmen *et al.* (2005, 2011) relate—for a subset of 450 individuals—the answers to the behavior in a field experiment with a monetary lottery and find a strong correlation. Dohmen *et al.* (2005) also report on a strong correlation between the risk-scale answers and the answers to hypothetical lottery questions.

[4] For the history of the MPL instrument and other instruments, refer to the survey by Harrison

the MPL, the subjects are asked to choose among two lotteries, a "safe" lottery A with a close pair of prizes, and a "risky" lottery B with prizes much further apart. The probability with which the high prize is selected (common for both lotteries) is then gradually increased as one moves down the list (from 0.1 to 1). The number of times the subject chooses the safe lottery is often used as a measure of her risk aversion.[5] The MPL experiments have been run both when prizes are hypothetical and when they are real.[6] Holt and Laury (2002) found that the subjects did not display significant difference in behaviour when the prizes where low and real (2 and 1.6 USD for the A lottery, 3.85 and .1 USD for the B lottery), or when they where scaled up (20, 50, 90 times) but hypothetical. However, they also found that subjects display significantly more risk averse behaviour when these real prizes are scaled up with a factor 20 than when real lottery prizes are low. This result was confirmed by Holt and Laury (2005) in a follow-up paper where they produced new evidence that upward scaling of real lottery prizes leads subjects to make more 'risk averse' choices, while similar scaling of hypothetical lottery prizes has no such effect.[7] Thus a careful conclusion of the experimental literature goes into the direction of identifying a discrepancy between revealed and stated preferences.[8]

We are interested in risk attitudes towards large income risks in developed OECD countries. Elicitation of such attitudes by means of real prize lotteries is difficult. First, such experiments would be prohibitively costly.[9] Second, even

and Rutström (2008, section 1).

[5]A more rigorous measure is obtained by assuming a particular form for the subjects's von Neuman-Morgenstern (vNM) utility function (often the CRRA function), and estimating the risk aversion parameter that maximises the likelihood of the observed choices.

[6]Often, the runs with real prizes make use of a random problem selection procedure: out of the ten pairs of lotteries, a single pair is drawn at random and played out. The subject is then paid out according to the choice that it made for that lottery.

[7]Harrison *et al.* (2005) drew attention to the fact that the Holt and Laury (2002) conclusions are confounded by order effects since the subjects' experience with playing the 'low real' MPL could affect their behaviour in the subsequent 'high real' runs. The experiments they run (a between-subjects analysis for identifying the order effect, and a within-subjects analaysis for identifying the scale effects) show that there is indeed a scale effect but that the size estimated by Holt and Laury (2002) is upward biased. The Holt and Laury (2005) paper isolates the scale effects.

[8]Some studies, like Noussair *et al.* (2014) find no significant differences for any of the attitudes towards risk under scrutiny (risk aversion, prudence, temperance). However, their conclusions are based on a comparison of average choices, i.e., on a "between subjects"-analysis not controlling for individual heterogeneity, and this may hide some of the effects. E.g., when reexamining the data from Battalio *et al.* (1990) by means of a "within subjects"-analysis, Harrison (2005) finds that the conclusion of that study (which was based on a between-subjects analysis) is turned into a significant increase in difference in risk attitudes elicited in a real prize experiment and those in a hypothetical prize experiment.

[9]For this reason, researchers have run real prize lottery experiments in low income countries,

if budgets for significant positive prizes are available, it would not be possible to expose subjects to significant losses–unless they were informed in advance about the risk of loosing a large amount of their income, in which case the experiment would be prone to a strong self-selection effect.[10] Even for experiments with small or moderate lottery prizes, sample selection bias is considered to be a real problem. As stated by Harrison and Rutström in their extensive survey of the experimental literature: "All that is required for sample selection to introduce a bias in the risk attitude of participants is the expectation of uncertainty, not the actual presence of uncertainty in the experimental task" (Harrison and Rutström, 2008: 125). E.g., the results of Harrison *et al.* (2009) indicate that measured risk aversion is smaller when corrected for sample selection, probably due to fact that the announcement of a guaranteed show-up fee encourages subjects with an above-average risk aversion to participate.[11]

The unavailability of real lotteries with large stakes leaves hypothetical income gambles as the alternative instrument.[12] There are arguments for why the risk of bias when using this alternative instrument is limited. First, the instrument's hypothetical nature essentially removes selection on the endogenous variable–risk aversion. Second, hypothetical lotteries with large stakes may force respondents to provide a more informed choice answer then when choosing among lotteries (hypothetical or real) with small stakes; in other words, the proneness to make errors may be smaller (but not necessarily absent) in the former case.[13] Third, studies that have used the BJKS instrument or a similar one find that the answers

exploiting the fact that the prizes which the research budget allows for are large relative to the average monthly incomes in some groups of society. See, e.g., Kachelmeier and Shehata (1992) with students from Beijing University as subjects.

[10] In lab experiments on risk aversion, the size of losses is bounded from above by the participation fee such that no subject risks leaving the experiment with less money than when entering. This means that the loss is 'framed'–it is not a loss w.r.t. the income prior to joining the experiment (Harrison and Ruthström, 2008: 73).

[11] Sample selection bias may go in two directions: (i) the knowledge or belief of randomisation in the experiment can attract individuals that are more risk tolerant than the population at large; (ii) the show-up fee can encourage individuals that are more risk averse than the population at large to participate. The results of Harrison *et al.* (2009) indicate that the second bias dominates in their lab/field experiment.

[12] An anonymuous referee suggested a hybrid solution that relies on estimating a bias function (Blackburn *et al.*, 1994). In the present setting, this would consist of asking subjects to make choices from menus of both real and hypothetical low scale lotteries. Choice behaviour is then estimated conditional on respondents' observable characteristics. Next, the probability of choices from the low scale real lottery is computed conditional upon characteristics and the choice from the low scale hypothetical lottery. Assuming stability of this conditional probability in the stake size, it is finally applied to the choices from high stake hypothetical lotteries.

[13] Relatedly, the BJKS lotteries are fairly simple: subjects are asked to state their choice intentions between a sure income and a binary income lottery. I.e., they need not state a certainty equivalent or a probability premium.

to the hypothetical lottery questions or measures of risk attitudes derived from those answers correlate well with different kinds of risk behaviours, such as financial risk taking, occupational choice and health behaviours: Kimball *et al.* (2008) and Sahm (2012) find a strong association of elicited risk tolerance on stock holding; Ahn (2010) and Brown *et al.* (2011) find strong associations with the decision to become self-employed; Falk *et al.* (2017) find strong correlation between the risk tolerance measure derived from the answers to the hypothetical "staircase risk task" and self-employment and smoking intensity.[14]

Irrespective of whether stakes are real of hypothetical, there is still the issue that participation into the survey may be prone to selection because people with a larger opportunity cost of time may choose not to participate.[15] If all aspects of opportunity costs are controlled for, and unobservable determinants of risk aversion are otherwise uncorrelated with those of the participation decision, then selection on observables will not lead to inconsistent estimates of the risk aversion equation parameters. However, to compute an unbiased estimate of the average risk aversion in the population, it will be necessary to correct the sample estimate for selection on observables–e.g., to correct for the fact that underrepresentation of higher income groups biases the population average upwards (if such groups are less risk averse); we will come back to this issue at the end of Section 3.

2.2 The BJKS hypothetical income lotteries

We use the BJKS instrument to measure risk attitudes in Norway, thus extending the existing sample of five countries with a sixth one that is known to have a large welfare state. We will estimate a stated risk aversion measure and argue that it correlates well with various risk behaviours. We then use it to compute an average degree of relative risk aversion that will be compared with that for the other countries.

The BJKS question reads as follows:[16]

[14] Another hypothetical instrument is the answer to the question asked in the Bank of Italy's 1995 SWIH survey. Respondents were asked about their willingness to pay for a security paying off 10 mln Lire (about 5000 Euro) or nothing at all, both with probability $\frac{1}{2}$. Based on the respondent's answer, Guiso and Paiella (2006) calculate the implied coefficient of absolute risk aversion and show that it has predictive power for decisions in the domain of financial investment (both ownership and share of risky financial assets in the household's portfolio) and occupational choice. Using Dutch data, Cramer *et al.* (2002) find a strong negative correlation between the WTP answers for a hypothetical lottery and self-employment status on the labour market.

[15] An example for such selection in an experiment is given in Rutström (1998).

[16] Kimball *et al.* (2008) argue that the original question used by BJKS gives rise to a *status quo* bias because it asks the respondent to choose between his/her current job (no change) and a new job with a risky income prospect. A biased answer may then be expected because the respondent is likely to put weight on the known non-monetary aspects of the current job (the

"Suppose that you are the only income earner in your household. Suppose also that reasons beyond your control force you to change occupation. You can choose between two alternatives. Job 1 guarantees you the same income as your current income. Job 2 gives you a 50% chance of an income twice as high as your current income, but with a 50% chance it results in a reduction of your current income by one third. What is your immediate reaction? Would you choose job 1 or job 2?"

If the respondent selects the safe alternative (job 1), she is presented with a new pair of alternatives, the only difference being that the downside risk of job 2 is one fifth of the current income instead of one third. If, on the other hand, job 2 is selected, a follow-up question presents the respondent with a choice between the safe alternative and a risky job 2 where the downside risk is increased from one third to one half.

Suppose that individual preferences over income lotteries can be represented by a continuous function $V(\cdot)$. Let λ be defined as the scaling factor that makes an individual indifferent between the lottery $(\lambda C, 2C; \frac{1}{2}, \frac{1}{2})$ and the sure outcome lottery $(C, C; \frac{1}{2}, \frac{1}{2})$, i.e.,

$$V(\lambda C, 2C; \frac{1}{2}, \frac{1}{2}) = V(C, C; \frac{1}{2}, \frac{1}{2}). \tag{1}$$

With monotone preferences, we can infer from the way the respondent answers the lottery questions, to which of the following intervals her λ belongs: $[0, \frac{1}{2}]$, $(\frac{1}{2}, \frac{2}{3}]$, $(\frac{2}{3}, \frac{4}{5}]$, or $(\frac{4}{5}, 1]$. Even at this level of generality, it is natural to think of λ as a unit-free measure of aversion towards risk taking.[17]

working environment, the nature of the job tasks, etc.) (cf Samuelson and Zeckhauser, 1988). To avoid this possibility of bias, they propose a framing that is not linked to the current situation. We follow this advice and rather than asking respondents about a choice between their current job and a risky alternative, as in the original BJKS formulation, we ask respondents to choose between two new jobs, one with a certain and another with a risky income prospect.The framing of this question is different from the one used in Kimball *et al.* (2008) and Kapteyn and Teppa (2011) who seek to remove the status quo bias by suggesting that the cause of job change was rooted in an allergy problem. Instead, we use the phrase "factors beyond your control" so as not to make the decision to rely on a specific disease/problem that some respondents might deem as a remote cause for a job change.

[17] $\lambda = 1$ corresponds to extreme risk aversion, while $\lambda = 0$ corresponds to risk neutrality. If preferences admit an expected utility representation with von Neumann Morgenstern utility function $u(\cdot)$, the former case would turn $u(\cdot)$ flat for any income above today's income (i.e., extreme risk aversion), while $\lambda = 0$ corresponds to $u(\cdot)$ being linear. Global risk loving behavior is thus ruled out, because it would require negative income levels. Since income in the lottery questions has the flavour of permanent income, this non-negativity requirement (and therefore the exclusion of global risk loving) stands to reason. Even though the utility function may display locally convex parts to begin with, the availability of fair or almost fair gambling opportunities would result in a concavification of such convex parts (see Raiffa, 1968, pp. 94-96, or Drèze, 1971, sec. 2.1).

3 The distribution of stated risk preferences in Norway

3.1 The 2006 survey

The data for our study were gathered through a survey conducted by a major market intelligence company (Synovate–now Ipsos) in the spring of 2006. It was commissioned by a large Nordic insurance carrier as part of a study on people's attitudes to issues of safety, security, anxiety, trust, etc. The target group consisted of people in the age group 18–74. The response rate was 57.4% with 1554 responses.[18] In the survey, people were taken first through a list of 26 questions, asking what makes them feel safe and secure and which situations they fear most, as well as inquiring about trust and to which extent they feel satisfied with their current state of life. Next, they were asked to consider the hypothetical income gambles, previously described. Finally, they were asked about socioeconomic characteristics. See Table A1 in Appendix A for descriptive statistics.

Table 1 gives information about responses to the income lottery gambles. More than 75 percent expressed a choice intention for low or moderate risk. From the cumulative frequencies it is clear that the stated risk aversion distribution for women first order stochastically dominates that for men. On the other hand, the distributions for higher age, lower educational attainment, and lower income level only second order stochastically dominate those for lower age, higher educational attainment and higher income. Thus a more careful econometric analysis is needed to shed light on these relationships.

3.2 Estimation of the stated risk attitude distribution

We now explore the relationship between the elicited (stated) attitude toward income risk and socioeconomic characteristics. Rather than making any assumption on a particular type of Bernoulli utility function (or whether preferences have a vNM-representation, for that matter), we perform our econometric analysis on λ which was shown to belong to one of the following four intervals: $[0, \frac{1}{2}]$, $(\frac{1}{2}, \frac{2}{3}]$, $(\frac{2}{3}, \frac{4}{5}]$, or $(\frac{4}{5}, 1]$. Because λ_i itself is not observable, we regard it as a latent variable

[18] For the age group 18–54, the sample was taken from a representative e-base while older respondents were randomly drawn from a representative postal base. The reason for using standard mail for the latter group is that penetration of the Internet declines with age. Internet use at home in 2006 for different age groups in Norway is as follows: 16-24: 92%, 25-34: 86%, 35-44: 80%, 45-54: 76%, 55-64:57%, 65-74: 28% (Statistics Norway, *Statistikkbanken*, Table 07002). Both the e-base and postal base are built upon the national telephone directory. The response rate for those who answered the questionnaire on the Internet was 56.6 percent. For the postal survey, the response rate was 59.4 percent.

Table 1: (Cumulative) distribution of responses to the income lotteries

	(1)	(2)	(3)	(4)
	"Job 2","Job 2" $\lambda<\frac{1}{2}$	"Job 2","Job 1" $\frac{1}{2}<\lambda<\frac{2}{3}$	"Job 1","Job 2" $\frac{2}{3}<\lambda<\frac{4}{5}$	"Job 1","Job 1" $\frac{4}{5}<\lambda<1$
All	13.31	8.59(21.9)	41.31(63.21)	36.79(100)
Men	13.66	10.40(24.00)	41.16(65.22)	34.78(100)
Women	8.45	4.08(12.53)	49.03(61.56)	38.44(100)
18-34 y.o.	19.90	11.01(30.91)	28.61(59.52)	40.48(100)
34-49 y.o.	12.32	9.98(22.30)	39.65(61.95)	38.14(100)
50-64 y.o.	7.24	5.35(12.59)	51.95(64.18)	35.47(100)
65+y.o.	6.30	2.10(8.4)	60.57(68.97)	31.04(100)
Primary school	4.43	3.78(8.21)	73.45(81.66)	18.34(100)
Secondary school	9.06	7.27(16.33)	47.51(63.84)	36.17(100)
University	13.18	7.64(20.82)	40.80(61.62)	38.39(100)
<300k NOK[a]	10.66	5.74(16.40)	47.76(64.16)	35.85(100)
300-500k NOK	10.84	6.62(17.46)	45.61(63.07)	36.92(100)
>500k NOK	12.82	12.60(25.42)	37.14(62.56)	37.44(100)
Student	21.20	11.60(32.80)	27.62(60.42)	39.57(100)
Employed	11.76	8.52(20.28)	42.04(62.32)	37.68(100)
Unempl/retired	7.14	2.91(10.05)	57.25(67.30)	32.71(100)

[a] 1NOK=.093EUR=.112USD (2006, PPP)

depending on the vector of observable co-variates x_i and a random component ε_i. We assume that

$$\log \frac{\lambda_i}{1-\lambda_i} = x_i'\beta + \varepsilon_i,\ \varepsilon_i \sim N(0,\sigma^2). \tag{2}$$

The left-hand variable is latent, but we observe whether it belongs to the interval $(-\infty, 0]$, $(0, \log 2]$, $(\log 2, \log 4]$, or $(\log 4, \infty)$. Thus we have interval-coded data and estimate β and σ using interval regression (Wooldridge, 2010, ch 19.2.2).[19]

For the purpose of the international comparison in Section 4, we first estimate (2) without any covariates, i.e., with only a constant term β_0. The result is given in column (1) of Table 2. To interpret the result, we calculate the conditional expectation $\widehat{\lambda}_i = E(\lambda_i|\widehat{\beta}_0, \widehat{\sigma}, L_i)$ where $L_i \in \{[0, \frac{1}{2}], (\frac{1}{2}, \frac{2}{3}], (\frac{2}{3}, \frac{4}{5}], (\frac{4}{5}, 1]\}$ indicates the interval that respondent i has implicitly chosen for λ_i (see appendix C for details). The conditional and unconditional means are given in column (1) of the upper panel of Table 3. The average respondent has an expected λ of .731, and since the median respondent belongs to $(\frac{2}{3}, \frac{4}{5}]$, the median λ is .738.

The estimation results for the conditional model are given in column (2) of Table 2. The descriptive statistics for the co-variates and others employed later in the paper are given in Appendix A. *A priori*, there is no guarantee that these estimates are unbiased, because survey non-response may not be completely at random. Some of the non-response may be due to selection on observables. In this case, it can be shown that unweighted conditional maximum likelihood estimation is consistent (cf Wooldridge, 2002, ch 19.8). If sample selection also happens on unobservables, that is on the risk aversion of the respondent, then consistent inference must also take into account the selection process, requiring the formulation and estimation of a model of how subjects select into the sample. The earlier mentioned study by Harrison *et al.* (2009) diagnoses such selection on the variable of interest and therefore on unobservables. Since we do not observe anything on the non-responding subjects we cannot test for selection on unobservables. There are two arguments for why we should not expect such selection to be important: the income lotteries were hypothetical—nothing could be gained (or lost) by participating—and the survey was not of a nature that might trigger any impression of being able to make a case (e.g., as in surveys on political or environmental issues). E.g., in a recent study on risk preferences in The Netherlands, von Gaudecker *et al.* (2012) find little/no evidence of self-selection due to non-participation/incomplete participation within a randomly drawn sample from the broad population.

In line with many other studies, men are more risk tolerant than women.[20]

[19] Thus we assume λ_i follows a logit-normal distribution since it is the logit transform of a normally distributed variable. The flexibility of the logit-normal is shown in Lesaffre *et al.* (2007) and Andersen *et al.* (2012).

[20] See, e.g., the review by Croson and Gneezy 2009). See Filippin and Crosetto (forthcoming)

Table 2: Ordered probit estimation of equation (4)

	(1)	(2)	(3)
	ML model (2)		Ordered probit
constant	1.150*** (.0243)	1.229*** (.2489)	\
age	\	.0012 (.0031)	.0006 (.0037)
male	\	−.6949*** (.1600)	−.8090*** (.1910)
age×male	\	.0107*** (.0033)	.0125*** (.0039)
cut off$_1$	0	0	−1.253 (.2987)
cut off$_2$	log 2	log 2	−.8969 (.2984)
cut off$_3$	log 4	log 4	.2563 (.2978)
σ	.8847*** (.0302)	.8449*** (.0288)	1
McKelvey & Zavoina's R^2	0	0.057	0.071
Log lik	-1935.01	-1886.32	-1784.91
N	1509	1509	1509

[a] Both models were estimated with the following controls: dummies for educational attainment, civil status, labour market status, income scale, county of residence, type of residential area, life satisfaction, religiosity. Robust standard errors. Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test). Wald test statistic for H_0 :"all income dummies are zero": $\Pr\{\chi^2(8) > 3.13) = .9258$ (interval regression) and $\Pr\{\chi^2(8) > 3.56) = .8944$ (ordered probit).

Age does not correlate with stated risk aversion, unless interacted with the gender dummy: the choice intentions of elderly men in our sample are less 'risky'. However, since our sample is a cross section, the age variable may also pick up cohort effects. These results are in line with previous research, with some minor qualifications.[21] None of the eight income scale dummies are significant, and the hypothesis that all dummies are jointly zero cannot be rejected. Like in Dohmen *et al* (2011) we find significant positive correlations (not shown) between stated risk aversion and both life satisfaction and religiosity. Both control variables, however, are likely to be correlated with the error term and the estimated coefficients can thus not be interpreted as measures of a causal effect (see supplementary appendix J for details). Column (3) of Table 2 shows that these conclusions are preserved when estimating an ordered probit model for the four risk groups. Table G1 in supplementary appendix G shows the robustness of the interval regression results by sequentially introducing the different sets of controls.

We next calculate for each individual in the sample the conditional expectation $\widehat{\lambda}_i = E(\lambda_i | x_i'\widehat{\beta}, \widehat{\sigma}, L_i)$. The descriptive statistics for the imputed λ's are given in columns (2)-(6) of the upper panel Table 3. The means reported in column (2) are identical to the unconditional means in column (1). Column (4) reports on the sample standard deviation; it ignores the fact that there is a variance around each $E(\lambda_i | x_i'\widehat{\beta}, \widehat{\sigma}, L_i)$ due to $\widehat{\sigma} > 0$.[22] Most of the heterogeneity in stated risk aversion is inter-group. In appendix D, we validate $\widehat{\lambda}_i$ as a measure of stated risk attitude by relating it to various instances of risk behaviour. We have estimated probit models for daily smoking, being obese (BMI>30), working in the private sector, having a top manager position, as well as an ordered probit model for the stated likelihood of stock investment with borrowed funds. In all five cases, our stated risk aversion measure has a significant negative correlation with the (stated) risk behaviour.

If we restrict $V(\cdot)$ to an expected utility form and assume that the vNM utility function $u(\cdot)$ has constant relative risk aversion R, then a monotone positive

for a meta study with counter-evidence.

[21] Strong gender and age effects are found in BJKS, in Halek and Eisenhauer (2001), Dohmen *et al.* (2011), Sahm (2012), Dohmen *et al.* (2017) and Falk *et al.* (2017). BJKS find non-monotone age effects (for respondents age 51 or higher). Guiso and Paiella (2008) also document age effects, but do not find a strong gender effect.

[22] The values for $E_{i \in L} var(\lambda | x_i'\widehat{\beta}, \widehat{\sigma}, L)$ are .006258, .0022165, .001446 and .001959 for $L = [0, \frac{1}{2}], (\frac{1}{2}, \frac{2}{3}], (\frac{2}{3}, \frac{4}{5}]$ and $(\frac{4}{5}, 1]$, respectively. For the unconditional model, the sample standard deviations are by definition zero, but the variance around the mean is larger (due to a larger $\widehat{\sigma}$).

Table 3: Conditional and unconditional sample distributions of $\widehat{\lambda}_i and \widehat{R}_i$.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	mean[a]	mean[b]	median[b]	st.dev.[b]	min[b]	max[b]	N
$\widehat{\lambda}_i\|_{L_i=(0,\frac{1}{2})}$	.401	.401	.403	.011	.356	.421	201
$\widehat{\lambda}_i\|_{L_i=(\frac{1}{2},\frac{2}{3})}$	.595	.594	.594	.003	.587	.602	130
$\widehat{\lambda}_i\|_{L_i=(\frac{2}{3},\frac{4}{5})}$	.738	.738	.739	.002	.731	.744	621
$\widehat{\lambda}_i\|_{L_i=(\frac{4}{5},1)}$	.873	.872	.872	.006	.855	.890	557
$\widehat{\lambda}_i$	.731[c]	.730[c]	.740	.153	.356	.890	1509
$\widehat{R}_i\|_{L_i=(0,\frac{1}{2})}$	.658	.656	.662	.031	.535	.713	201
$\widehat{R}_i\|_{L_i=(\frac{1}{2},\frac{2}{3})}$	1.527	1.522	1.521	.019	1.477	1.573	130
$\widehat{R}_i\|_{L_i=(\frac{2}{3},\frac{4}{5})}$	2.833	2.838	2.841	.029	2.737	2.914	621
$\widehat{R}_i\|_{L_i=(\frac{4}{5},1)}$	6.734	6.667	6.663	.419	5.648	7.980	557
$\widehat{R}_i$	3.867[c]	3.847[c]	2.854	2.290	.535	7.980	1509

Source: own calculations

[a] mean based on column (1) of Table 2. [b] statistics based on column (2) of Table 2. [c] weighted average of the four preceding figures in same column, weights are the fractions of respondents in each response category.

relationship between λ and R follows (see BJKS):[23]

$$\lambda = \left[2 - 2^{1-R}\right]^{\frac{1}{1-R}} . \tag{3}$$

The mapping from λ to R is shown in Figure 1. The corresponding intervals for R_i are $[0, 1]$, $(1, 2]$, $(2, 3.76]$, and $(3.76, \infty)$. The lower panel of Table 3 give the descriptive statistics for the imputed $\widehat{R}_i$ (see appendix C for details on computation). The sample average for $\widehat{R}$ is 3.85 with a standard deviation of 2.29. Without controlling for observable characteristics we get a marginally higher value (3.87).

Even without sample selection on unobservables, the estimate of the mean risk aversion measure for the population may be biased if selection happens on observables. In supplementary appendix F, a comparison of the composition of our sample with the Norwegian population shows that some groups are under/over-represented. However, we also show there that when accounting for this under/over-representation by applying the Bethlehem and Keller (1987) linear weighting scheme, the sample averages for $\widehat{\lambda}$ and $\widehat{R}$ remain virtually identical. Thus we feel confident that the sample averages are representative for the entire population.

[23] Without imposing a specific functional form, but using standard Taylor expansions around the mean, we get $-\frac{u''(E\widetilde{C})}{u'(E\widetilde{C})}E\widetilde{C} \simeq \lambda\left(1+\frac{\lambda}{2}\right)\Big/\left(1-\frac{\lambda}{2}\right)^2$. However, the risks are too large to justify this local measure of relative risk aversion.

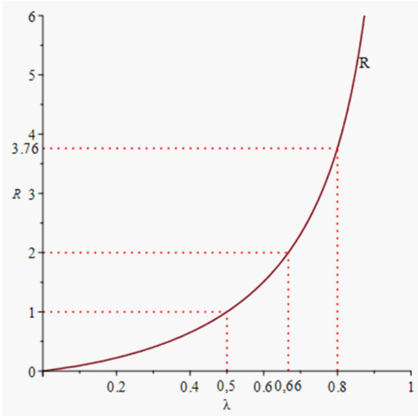

Figure 1. The mapping from λ to R.

4 International comparison

4.1 The role of background risks

The multiplicative nature of the risk in the BJKS lottery questions makes the elicited risk aversion measures unit free and therefore comparable both over time and across countries. In Table 4, columns (1)-(4), we have reproduced the answer distribution to the BJKS lottery questions for Norway and five other surveys, all collected before the financial crisis of 2008: USA 2002 ($N = 3591$), France 2004 ($N = 3674$), Chile 2002 ($N = 11475$), The Netherlands 2004 ($N = 1315$), and Italy 2007 ($N = 1686$).[24]

On the basis of these response distributions we have estimated for each country a mean and standard deviation for λ and a mean for the (constant) coefficient of relative risk aversion in the same way as was done for Norway in column (1) of Table 3.[25] The results are displayed in columns (5)-(7) of Table 4. For the US, the value is half a unit above the one that Kimball *et al.* (2008, Table 4) report when

[24] US: Kimball *et al.* (2008), Health and Retirement Survey (HRS); NL: Kapteyn and Teppa (2011), CentER internet panel, FR: Arrondel and Savignac (2015) (INSEE Wealth Survey); CL: Martinez & Sahm (2008), Chilean Social Security Survey; IT: Butler *et al.* (2011), Unicredit Clients' Survey. We are grateful to Véronique Flambard (Université catholique de Lille) for drawing our attention to the 2004 INSEE Wealth Survey.

[25] It therefore suffices to have information on the number of respondents that end up in the different risk aversion classes. The ML estimates for Norway where given in column (1) of Table 2. The estimates for the remaining countries are: FR 2004: $\widehat{\beta}_0 = 1.589$ (.021), $\widehat{\sigma} = .901$ (.021), NL 2004: $\widehat{\beta}_0 = 1.509$ (.041), $\widehat{\sigma} = 1.12$ (.043), US 2002: $\widehat{\beta}_0 = 1.947$ (.043), $\widehat{\sigma} = 1.620$ (.048), CL 2002: $\widehat{\beta}_0 = 2.987$ (.049), $\widehat{\sigma} = 1.756$ (.040).

Table 4: Cross-country comparisons of responses to the income gamble.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	$0 < \lambda < \frac{1}{2}$	$\frac{1}{2} < \lambda < \frac{2}{3}$	$\frac{2}{3} < \lambda < \frac{4}{5}$	$\frac{4}{5} < \lambda < 1$	$E(\lambda)$	$\sigma(\lambda)$	$E(R)$
Chile 2002[a]	4.7	4.1	9.3	81.8	.89	.16	10.52
USA 2002[a]	11.7	9.6	15.3	63.4	.80	.21	8.73
France 2004	4.7	10.3	26.5	58.5	.80	.14	5.55
Netherlands 2004[a]	7.9	17.8	19.4	54.9	.77	.17	5.73
Norway 2006	13.3	8.6	41.3	36.8	.73	.16	3.87
Italy 2007	13.2	8.0	26.1	53.7	.76	.19	5.97

[a] In these surveys, respondents choosing job 2 in both rounds were given the choice in a third round between job 1 and a job with a 50% risk of receiving only $\frac{1}{4}$ of current income. Likewise, respondents choosing job 1 in both rounds were given the choice in a third round between job 1 and a job with a 50% risk of receiving $\frac{9}{10}$ of current income. Respondents thereby sort themselves into 6 risk aversion classes. The class $\lambda <\frac{1}{2}$ is the sum of the classes $\lambda <\frac{1}{4}$ and $\frac{1}{4}< \lambda <\frac{1}{2}$. Likewise, the class $\frac{4}{5}< \lambda < 1$ is the sum of the classes $\frac{4}{5}< \lambda <\frac{9}{10}$ and $\frac{9}{10}< \lambda < 1$.

making use of the 1992 and 1994 waves of the HRS, and correcting for response errors (8.2).

Mean stated relative risk aversion is lowest in Norway. The middle position taken by France, The Netherlands and Italy is followed by the US, while Chile has the highest value. There may be three possible reasons for this variation across countries. First, it could stem from differences in sample composition. Should, say, the US sample consist of a much larger share of women than the Norwegian one and women are much more risk averse than men, then this could result in a twice as large value for $E(R)$. We do not believe such composition effects are very important, though: while there may be differences in the distributions of age and other variables that affect stated risk aversion, our analysis in section 3 and that of Sahm (2012) showed that these effects, while statistically significant, are modest. A second reason for international heterogeneity could simply be due to differences in the nature or nurture of risk preferences: Norwegians may just be born as risk tolerant or they are primed that life in their country is safe and stable. A third reason could be differences in social and economic context. Even though within each country stated risk attitudes may be fairly insensitive to income (as it is for Norway–cf the discussion of Table 3), average Chilean real income is of a lower order of magnitude than that for the other countries, and this may explain the two digit level for average stated relative risk aversion in that country. On the other hand, countries like Norway and the US with GDP per capita figures of the same order of magnitude have different levels of social protection and leave citizens with different levels of background risk. Intuitively, one would expect that the presence of an undesirable background risk increases the aversion towards other independent risks To illustrate, a potentially significant source of background risk are medical expenditure shocks arising because of incomplete health insurance. For a person paying a fixed or risk-rated health insurance premium the prospect of experiencing a 20% income fall and at the same time receiving a medical co-payment bill of \$6000 may easily overshadow that of an income doubling.[26] This need not be the case if almost complete health insurance is available at an earnings-related premium. Thus less complete health insurance leaves citizens with a higher background risk, which induces higher risk aversion (In appendix B, we show that this conclusion certainly holds when an actuarially fair insurance scheme is made marginally less fair.).

In the theoretical literature on background risk, undesirable background risk ($\widetilde{\varepsilon}$) has been defined in three ways: (i) as an unfair risk ($E\widetilde{\varepsilon} \leq 0$), (ii) as a risk that decreases expected utility ($Eu(y+\widetilde{\varepsilon}) \leq u(y)$) and (iii) as a risk that increases expected marginal utility ($Eu'(y+\widetilde{\varepsilon}) \geq u'(y)$). The decision maker is then said to be vulnerable to background risk if its introduction never makes any undesirable foreground risk desirable.[27] Depending on the type of background risk considered risk preferences are said to exhibit (i) risk vulnerability (Gollier and Pratt, 1996), (ii) properness (Pratt and Zeckhauser, 1987) or (iii) standardness (Kimball, 1993). All three notions of vulnerability imply that the decision maker becomes more risk averse when undesirable background risk is introduced. Gollier and Pratt (1996) and Gollier (2001, ch 9) show that standardness implies properness which in turn implies risk vulnerability which in turn implies decreasing absolute risk aversion (DARA–because $E\widetilde{\varepsilon} \leq 0$ covers the degenerate case of a certain negative value for ε). Necessary and sufficient conditions for standardness are DARA and the coefficient of absolute prudence ($-\frac{u'''}{u''}$) falling in income. Hence, these are also sufficient for (i) and (ii). Another set of sufficient conditions for (i) is that the coefficient for absolute risk aversion is falling and convex. These conditions are not trivial, but at the same time natural assumptions to make (risk aversion and DARA imply that absolute risk aversion must be predominantly convex in wealth). Decreasing absolute prudence means that the precautionary savings motive falls with wealth.[28] Beaud and Willinger (2015) present evidence from a within-subject

[26]Goldman and Maestas (2012, Table 3) report that in 1999-2000 5% (1%) of US retirees with only Medicare A&B coverage, experienced out-of-pocket expenses in excess of \$6367 (\$31751).

[27]I.e., if for any initial wealth y and any foreground risk $\widetilde{x}$ such that $Eu(y+\widetilde{x}) \leq u(y)$ it follows that $Eu(y+\widetilde{x}+\widetilde{\varepsilon}) \leq Eu(y+\widetilde{\varepsilon})$.

[28]Eeckhoudt *et al.* (1996) extend these results by looking at first and second order stochastic dominated deteriorations of existing background risk. The conditions for such shifts to make the decision maker more averse to foreground risk are stronger than for an introduction of a background risk.

experiment for risk vulnerability.

Ideally, we would like to have a time series of $E(R)$ for each country so that a panel data model could discriminate the country specific fixed effects from the influences of macroeconomic and welfare state conditions. Unfortunately, such data do not exist. So we will proceed with a descriptive comparison of several indices and measures of welfare state generosity and the risks against which the welfare state aims to protect. For each risk, we will propose a ranking of the six countries and conclude that these rankings correlate well with the rankings in terms of stated risk aversion.[29]

4.2 Welfare state generosity

We consider three background income risks: unemployment risk (u), sickness risk (s–the risk of loosing earnings because illness prevents going to work) and health expenditure risk (h). Table 5 shows for each risk public spending as a fraction of GDP. These figures are the most common measures of a country's welfare state (e.g., Garfinkel *et al.*, 2010: 40); they give a first indication of how much of the background risk is covered by the welfare state.

These public spending figures can be decomposed into a replacement rate (RR, the complement of a coinsurance rate) for eligible workers/citizens, a beneficiary rate (BR, the % of the population at risk receiving the benefit), the size of the risk (p, the probability of getting unemployed or ill), and the fraction of total income "at risk" (X)–the wage bill if everybody remains employed or nobody calls in sick, the total medical expenditure should everybody fall ill. Thus we can write

$$\frac{P_j}{Y} = RR_j \times BR_j \times p_j \times \frac{X_j}{Y}, \tag{4}$$

where P_j is public expenditure related to risk j ($j = u, s, h$) and Y is GDP.

These decompositions are approximate because the beneficiary rate underestimates the share of wages earned by eligible workers in $\frac{P_u}{Y}$ and $\frac{P_s}{Y}$, and the share

[29]Recent studies have indirectly tested for the background risk hypothesis by relating the degree of protection against medical expenditure to either the amount of precautionary savings or the riskiness of the wealth portfolio. Kapteyn and Panis (2005) compare savings behaviour after retirement in Italy, The Netherlands and the US and relate the stronger desire to hold bequeathable wealth by US citizens to the less generous public coverage of medical expenses in that country. Goldman and Maestas (2012) find that Medicare benificiaries in the US who have chosen a highly protective Medicare HMO policy are significantly more inclined to hold risky assets than those with moderately protective Medigap or employer supplemental health insurance, who in turn have a stronger inclination than those without any supplementary insurance. Atella *et al.* (2012) find clear evidence that in European countries with a publicly financed national health service (NHS), +50 citizens with poor health status are significantly more likely to hold risky assets than those in non-NHS countries.

Table 5: Public expenditures related to three risks as a percentage of GDP.

	CL 2002	US 2002	FR 2004	NL 2004	NO[a] 2006	IT 2007
unemployment (%)	0.022	0.493	1.645	1.530	0.390	0.337
sick leave (%)	0.127[b]	0.227	0.616	1.770	3.065	0.470
health care (%)[c]	3.530[d]	4.491	4.700	3.075	4.547	5.161[e]
sum	3.679	5.211	6.961	6.376	8.002	5.967
GDP per capita[f]	11075	41284	30215	36244	51156	30449

[a] mainland GDP; [b] 2003-figure; [c] public expenditure on curative and rehabilitative care (crc); [d] public expenditure on current health care (incl. crc); [e] public expenditure on crc, services of long term nursing care and ancillary services to health care; [f] 2005 USD (PPP). Source: OECD.Stat, except for medical expenditure for Chile: Bitrán and Urcullo (2008: 102).

of health expenditure spent on eligible citizens in $\frac{P_h}{Y}$. On top, they hide a lot of heterogeneity within each country. Still, they are useful because they highlight the different sources for the welfare state's size (Atkinson, 1995). The replacement rate is the fraction of income loss compensated in case of sickness or unemployment, and the fraction of medical expenditures reimbursed in case of illness. A second measure of generosity is the beneficiary rate, which can be written as the product of the coverage rate, CR_j, and the take up rate, TR_j. The former measures the percentage of the population that is in principle entitled to the benefit through contribution, while the latter adjusts for the length of the qualification period, the duration of the income replacement as well as number of waiting days in case of unemployment or sickness leave.[30] Of course, moral hazard may make the size of the risk dependent on the generosity of protection against it, so p_j may depend on RR_j and TR_j. The last term in (4) denotes the fraction of total income "at risk" and reflects the importance of protection. To put this starkly: if a country mainly consists of capital owners then the wage bill when all workers are at work will be very modest and a generous unemployment or paid sickness leave scheme is of minor importance.

In the next sub-section, we will present data from various sources on the different sizes of risk and welfare state generosity, and relate these to our average stated risk aversion measure.

[30] At the level of the individual, the take up rate may fall short of one if beneficiaries voluntarily refrain from applying for the benefit (e.g., because of stigma) or it may exceed one if–for statistical reasons–recorded benefits include more than the contingent benefit (e.g., because unemployment benefits and unemployment assistance are statistically aggregated).

Table 6: Generosity indices, coverage rates and risk factors.

		CL 2002	US 2002	FR 2004	NL 2004	NO 2006	IT 2007
1	*Unemployment risk*						
1.1	CWED unemployment generosity index	4.4	10.3	11.3	12.0	14.0	11.1
1.2	CWED UB replacement rate (%)	6.0	58.4	70.6	78.7	66.5	59.6
1.3	CWED UB coverage rate (%)	58	89	87	88	91	88
1.4	Vroman unemployment generosity index (%)	0.9	14.3	28.0	25.0	21.4	13.2
1.5	Harmonised unemployment rate (%)	9.8	5.8	8.9	5.7	3.3	6.1
2	*Sickness leave risk*						
2.1	CWED sickness generosity index	10.4	0	12.3	14	15.9	9.7
2.2	CWED sickness absence replacement rate (%)	100	0	63.5	78.8	100	77.0
2.3	CWED sickness coverage rate (%)	61	0	100	89	100	69
2.4	Paid sickness leave in case of 5-day flu (days)	5	0	1	3.5	5	1
2.5	Paid sickness leave in case of 50-day cancer treatment (days)	50	0	29	35	50	24
2.6	Vroman sickness generosity index (%)	16.0	22.1	29.6	75.9	72.2	45.4
2.7	Days absence from work due to illness	4.6	4.4	9.8	13.5	17.2	5.9
2.8	Sickness absence odds ratio	0.020	0.018	0.041	0.048	0.079	0.026
3	*Medical expenditure risk*						
3.1	Public share of expenditure on curative and rehabilitative care (%)	48	46.9	84.1	66.5	85.3	82.9
3.2	Public health insurance coverage rate (%)	67.1	24.1	99.9	62.5	100	99.9
3.3	Fraction of population w/o health insurance (%)	16.2	15.2	0.1	2.1	0	0
3.4	Mortality rate at age 40 per 1000	1.7	2.0	1.7	1.1	1.0	0.9
3.5	Mortality rate at age 60 per 1000	10.2	10.0	7.9	8.1	6.7	6.1

See supplementary appendix H for sources and definitions.

4.3 Rankings of residual background risk

To measure welfare state generosity as well as the risks it insures against for the six countries we mainly rely on two datasets: the OECD.Stat database (http://stats.oecd.org) and the Comparative Welfare Entitlement Dataset (CWED), compiled by Scruggs *et al.* (2014a). The latter provides welfare state generosity measures for the first two income loss risks, unemployment and sickness (as well as for pensions) for 27 OECD countries, but unfortunately not for Chile. Hence, we construct the Chilean CWED index values on the basis of the Social Security Throughout The World files compiled by the US Social Security Administration (2003) using the CWED-recipe (Scruggs *et al.* 2014b). A different generosity measure is the Vroman index (Vroman, 2002, 2003) which can be written as $V_j = RR_j \times BR_j$. The generosity and risk measures are given in Table 7.

The CWED unemployment generosity index is built up from the replacement rates for a 40 year old average wage worker (single–given by row (1.2)–and cohabiting with dependent spouse without earnings and two children aged 7 and 12), the benefit qualification period, the benefit duration, and the percentage of the labour force insured for unemployment risk (coverage rate, given in row (1.3))–see Scruggs *et al.* (2014a). It is a measure of the degree of protection against income loss due to unemployment risk offered by a country's social insurance system. Chile scores worst on this index, intermediate positions are taken by the US, Italy and France, while the Netherlands and Norway score best.

Unlike the Scruggs index, the Vroman unemployment generosity index is not based on institutional but on macroeconomic variables; it is defined as

$$V_u = \frac{\text{av weekly unemployment benefit per worker}}{\text{average weekly wage}} \times \frac{\text{av weekly no. of recipients}}{\text{av weekly no. of unemployed}}.$$

An important difference between this index and that of Scruggs *et al.* (2014) is that the latter only takes the coverage rate into account, while the former relies on the beneficiary rate–the product of the coverage rate and the take up rate. V_u can be computed as the ratio of total unemployment compensation, P_u, to the aggregate wage bill, W, times the odds ratio of remaining employed:[31]

$$V_u = \frac{P_u}{W} \times \frac{1 - p_u}{p_u}. \tag{5}$$

The Vroman and CWED indices tend to "agree" that generosity is lowest in Chile, followed by the US and Italy, followed by France, the Netherlands and Norway. When also taking into account the probability of unemployment, measured by the harmonised unemployment rate at the time of the survey, we propose the following ranking w.r.t. exposure to unemployment risk:

$$NO \sim_u NL \prec_u FR \sim_u IT \sim_u US \prec_u CL.$$

Insurance against the risk of lost earnings due to sickness absence from work is measured by the CWED sickness generosity index. Like the index for unemployment generosity, it is built up from institutional variables such as the replacement rate, the length of the qualification period, the duration of the benefit and the coverage (fraction of the workforce covered). The highest score is for Norway, followed by the Netherlands and France, followed by Chile and Italy, followed by the US. Chile has a high replacement rate, making up in the index for an incomplete coverage of the workforce. The absence of a national sickness programme is

[31] Use the definition $V_u \overset{\text{def}}{=} RR \times BR$ and invert (4) to get $V_u = \frac{P_u}{X_u}\frac{1}{p_u}$. Since $X_u = wN$ where w is the average annual wage per worker and N is the labour force, and since the aggregate wage bill is $W = wN(1 - p_u)$ with p_u the unemployment rate, (5) follows.

responsible for the zero index value for the US.[32] In terms of replacement rate, the generosity of the different programmes is confirmed by the number of FT equivalent paid sickness leave days in case of a 5-day absence (e.g., flu) or a 50-day absence (for cancer treatment, say) (rows 2.4-5). The French and Italian, and to some extent also the Dutch, programmes have large "deductibles". If p_s denotes the fraction of working days lost due to illness, we can write $X_s = \frac{W}{1-p_s}$ and compute a Vroman index of generosity as

$$V_s = \frac{P_s}{W} \times \frac{1-p_s}{p_s}. \tag{6}$$

Row (2.8) lists the odds ratio $\frac{p_s}{1-p_s}$. Applying (6) then gives the Vroman index as given in row (2.6). The implied ranking w.r.t. sickness leave background risk more or less confirms that of the CWED index. To infer the residual background risk we should account for the risk of loosing working days due to illness. This is largest in Norway ($p_s = .073$) and smallest in the US ($p_s = .018$). However, due to moral hazard this risk is not independent of the different generosities: a more generous paid sickness leave arrangement may invite people to call in sick more often, while a high unemployment rate may discourage ill people to report ill. Hence the Norwegian figure is almost certainly overstating the "true" risk of illness,[33] while that for Chile may be understating it. We therefore assume a uniform risk of illness and rank the countries w.r.t. exposure to the risk of illness only on the basis of the Vroman and CDEW indices:

$$NO \sim_s NL \prec_s FR \sim_s IT \sim_s CL \prec_s US.$$

We finally consider medical expenditure risk. Protection against this risk is not measured in the CWED framework. But a good measure should take into account the replacement rate of those covered and the degree of coverage. Since total health expenditure can be written as $THE = p_h \times X_h$ (X_h can be thought of as average health expenditure when ill), we can approximate the Vroman generosity index for health care as $\frac{P_h}{THE}$,[34] which is given in row (3.1) of Table 7. According to this index, Norway, France and Italy are ranked top, followed by the Netherlands, followed by Chile and the US. However, this aggregate measure

[32] Six US states did have a sickness pay programme in 2002: Rhode Island, California, New Jersey, New York, Hawaii and Puerto Rico.

[33] Markussen *et al.* (2011) show that the high overall Norwegian level of sickness absence hides an unequal distribution: as much as 10% of the workforce can be absent for more than 15% of the time. They find that the Norwegian sickness insurance system is extensively used to cover other risks (e.g., traumatic personal events related to the family) than just own diagnosed illnesses, and that when the 100% replacement rate is about to terminate after one year, the weekly recovery rate makes a 6-fold jump, suggesting a strong moral hazard effect.

[34] Let the total and public per capita values be *the* and *phe* and let N_c (N_n) denote the number

hides some of the institutional aspects of health insurance in the different countries. In supplementary appendix I, we give a short summary of each country's health insurance system, which shows that the above ranking should be qualified by the following remarks:

- In Chile 2002, the public health insurance scheme (*Fonasa*) acts as an insurer of last resort, providing any citizen that does not wish to opt-out by seeking private health insurance (PHI), coverage at a premium that is 7% of taxable income. If publicly insured, a patient can only access publicly provided health care for which long waiting lists may apply.[35] Even when insured through *Fonasa*, Chileans face a risk of becoming seriously deprived due to out-of-pocket payments: in 2000 for about 87% of the households in the first income quintile, such payments exceeded 15% of their income (World Bank, 2004: 53-54).

- In the Netherlands 2004, private health insurers who cater to citizens with an income above a threshold are obliged by law (the *WTZ* act) to offer private insurees (who meet certain eligibility conditions and whose PHI premium exceeds the age-related WTZ premium) the same standard benefit package as under the public scheme (obligatory for those with incomes below the threshold). This provision ensures that everyone in the Netherlands has access to basic health insurance.

- In Norway 2006, there is almost complete coverage against health expenditure risk, but it is accompanied by a waiting time risk for elective treatments.

- In Italy 2007, the national health system offers almost complete coverage, but there exists a waiting time risk, especially for diagnostics.

- In the US 2002, public health insurance is only available to retired and poor people. It still leaves a considerable amount of financial risk due to

of citizens covered (not covered) by public health insurance. Then

$$\frac{P_h}{Y} = \frac{phe}{the_c} \times \frac{N_c}{N} \times \beta \times \frac{THE}{Y} = RR_h \times BR_h \times \beta \times \frac{THE}{Y},$$

where $\beta \stackrel{\text{def}}{=} \frac{the_c}{(the_c N_c + the_n N_n)/N}$. It is *a priori* unclear how per capita health expenditure on a publicly covered citizen relates to that of the average citizen. $the_n N_n$ can be decomposed as expenditure by non-insured people and expenditure by privately insured ones. Uninsured persons (of which there is a significant fraction in CL and the US–cf row (3.3) of Table 6) are likely to have lower health expenditure; privately insured persons may have expenditure that exceeds that of publically covered persons. For $\beta \simeq 1$, $V_h \stackrel{\text{def}}{=} RR_h \times BR_h$ is about $\frac{P_h}{THE}$.

[35] *Fonasa* insurees may apply for a free-choice policy that allows them to use contracted private providers. Out-of-pocket payments for this policy are generally high (Holst *et al.*, 2004)

significant co-payments which are not capped. Citizens can insure against these co-payments through supplementary insurance at a premium that is not income related and that is subject to a medical history risk.

In the light of these remarks, we suggest that protection against medical expenditure risk is highest in Norway, France and Italy, followed by the Netherlands, followed by the US and Chile. The risk of being in need of health care is difficult to measure. A crude proxy is the mortality rate. For ages 40 and 60, it is given in rows (3.5) and (3.6) of Table 6. At age 60 the mortality rate is highest in Chile and the US, followed by France and the Netherlands, followed by Norway, followed by Italy. This suggests the following ranking w.r.t. exposure to medical expenditure risk:

$$NO \sim_h IT \sim_h FR \prec_h NL \prec_h US \sim_h CL.$$

This short survey of welfare state protection against three main risks clearly reveals two distinct levels of residual exposure to background risk: a higher American level (CL, US) and a lower European level (IT, FR, NL, NO), perhaps with Norway standing out with the lowest background risk. *Ceteris paribus*, risk vulnerability then implies that risk tolerance is larger in Europe than in the US and Chile, which is what Table 4 confirms.

Other things are of course not equal. The average living standard and the distribution around the average differ as well, which may impact in the average attitude towards risk. Countries also differ in the prevalence of informal insurance arrangements (through networks of relatives and friends) that may cushion at least partly against employment and health risks. At the level of individual households, such schemes may complement formal welfare states schemes. At a macro level, economic growth may be associated with a gradual substitution of formal insurance schemes for informal ones. Thus several factors confound the relation between welfare state protection and attitudes towards foreground risk.[36] Hence, our conjecture that welfare state protection against economic and health risks contributes to a lower risk aversion against 'foreground' risks needs further attention in future research.

[36] In this respect, Vieider *et al.* (2015) find that risk aversion as revealed in real payoff experiments conducted among university students in 30 countries is positively correlated with GDP *per capita* (p/c). Similarly, Rieger *et al.* (2015) using hypothetical lotteries among university students in 53 countries report on a positive correlation between average country risk aversion and GDP p/c. On possible explanation is imperfect substitution between informal and formal insurance schemes as countries develop. But there may also be a selection issue: if students' risk aversion depends negatively on student income (DARA), but average student income relative to GDP p/c depends sufficiently negative on GDP p/c (because a student's (or her parent's) resources matter less to gain access to higher education in developed countries). E.g., in Falk *et al.* (2017), using samples that are representative for the adult populations, the correlation between risk tolerance and GDP p/c is positive, small and insignificant.

5 Summary and conclusion

In this paper we first made use of the hypothetical income gamble questions developed by Barsky *et al.* (1997) to map the distribution of stated risk preferences in Norway 2006. We thus extended the small set of countries for which this instrument has been employed with a sixth country that is characterised by an extensive welfare state.

Under the assumption of risk preferences taking the expected utility form and having constant relative risk aversion, the stated risk attitude distribution for Norway translates into an average (median) relative risk aversion degree of 3.8 (2.9) with a sample standard deviation of 2.3. This sample average is well above the pivotal value if 1 (Eeckhoudt *et al.* 2005). At the same time, it is smaller than the corresponding values for three other European countries (between 5 and 6) and much smaller than those for the US and Chile (above 8). Our conjecture is that background risks loom less pronounced in more extensive welfare states; under risk vulnerability, this increases the tolerance w.r.t. income risks. The evidence that we collect on protection against unemployment, sickness leave and medical expenditure risk underscores this hypothesis. Still, further substantiation of this conjecture and related hypotheses, requires additional empirical work. It also requires international datasets on values and attitudes to be complemented with comparable measures of attitudes towards risk. In this regard, we would like to highlight the absence of any systematic data collection on comparable measures of attitudes towards risk at the level of the OECD or EU. Economic and financial risks permeate the decisions of consumers, workers, investors and managers. Measures for attitudes towards risks have been formalised since forty years and cheap, comparable and reliable instruments for eliciting these attitudes have been around for at least twenty years. They are crucial both for positive and normative analysis: to understand how agents make choices under risk and to evaluate the welfare consequences of insurance mechanisms. Still, none of the main European surveys (EU-SILC, EU-LFS, European Social Survey, European Values Study) nor the World Values Survey systematically include a quantitative and comparable measure on attitudes towards risk. We think it is timely to extend at least one of these surveys in this direction.

The relation between welfare state generosity and risk attitudes raises the question about the optimal size of a welfare state. This paper has been totally silent on this normative issue. But it is clear that moral hazard aspects will be part of the answer, as suggested by the perfect correlation between the CWED sickness generosity index and the average no. of days absent from work due to illness (rows 2.1 and 2.7 in Table 6).

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6 Appendices

A Descriptive statistics

Descriptive statistics are given in Table A1. For 75 respondents the income interval was imputed, as explained in supplementary appendix F.

Table A1. Descriptive statistics

	mean	st.dev.	min	max	N
man	.496	.50	0	1	1554
age	43.98	15.35	18	74	1554
income $\in (0, 300)$	.45	.50	0	1	1549
income $\in [300, 600)$	.45	.49	0	1	1549
income $\in [600, 800)$	.05	.37	0	1	1549
income $\in [800, \infty)$	.04	.19	0	1	1549
studying	.10	.29	0	1	1554
unemployed	.02	.14	0	1	1554
disability pension	.06	.23	0	1	1554
old age pension	.10	.30	0	1	1554
other reason for non LM particip	.03	.17	0	1	1554
Oslo area	.11	.32	0	1	1554
East	.40	.49	0	1	1554
South West	.10	.30	0	1	1554
West	.18	.38	0	1	1554
Mid	.12	.32	0	1	1554
North	.09	.29	0	1	1554
BMI	25.78	4.44	9.9	61.7	1550
smoker	.26	.44	0	1	1537
employed in private sector	.42	.49	0	1	1551
years of education ≤ 10	.13	.34	0	1	1548
years of education $\in [11, 13]$	.31	.46	0	1	1548
university short	.30	.46	0	1	1548
university long	.19	.39	0	1	1548
ongoing education	.06	.24	0	1	1548
country side	.31	.46	0	1	1553
small town (<20' inhabitants)	.16	.37	0	1	1553
medium town (20'-100' inhabitants)	.20	.40	0	1	1553
big town (>100' inhabitants)	.28	.47	0	1	1553
industry	.09	.08	0	1	1547
retail	.07	.07	0	1	1547

Table A1. Descriptive statistics: contd.

	mean	st.dev.	min	max	N
transport	.04	.04	0	1	1547
education/research	.13	.11	0	1	1547
health/social	.13	.11	0	1	1547
bank/insurance/finance	.02	.02	0	1	1547
energy	.01	.01	0	1	1547
IT/telecom	.04	.04	0	1	1547
other services	.06	.06	0	1	1547
other sector	.16	.13	0	1	1547
not working	.24	.18	0	1	1547
life satisfaction low[a]	.25	.44	0	1	1549
life satisfaction medium[a]	.58	.49	0	1	1549
life satisfaction high[a]	.17	.38	0	1	1549
religious? yes	.27	.20	0	1	1553
religious? no	.52	.25	0	1	1553
religious? do not know	.15	.13	0	1	1553
religious? don't want to answer	.05	.05	0	1	1553

[a] life satisfaction coded low/medium/high if answer 1-5/6-8/9.

B Incomplete insurance against a background risk leads to higher risk aversion

Suppose that a citizen with a vNM utility function $u(\cdot)$ and income y faces with probability p a loss L against which he can buy insurance a at a premium $(1+k)pa$. Let a^* be the optimal amount of insurance satisfying the first and second order conditions and $v(y;k) \stackrel{\text{def}}{=} (1-p)u(y-(1+k)pa^*) + pu(y-L+a^*-(1+k)pa^*)$ the corresponding maximal expected utility. The amount the agent is willing to pay for avoiding a small zero mean income risk ε, is given by the Arrow-Pratt risk premium which is proportional to the Arrow-Pratt coefficient of absolute risk aversion for $v(\cdot)$: $A^v = -\frac{v''(y;k)}{v'(y;k)}$. The question how this risk aversion coefficient changes with the loading factor k is non-trivial because in addition to giving a direct income effect on A^v, the increase in the loading factor will affect the demand for insurance (through substitution and income effects) which in turn affects A^v. A clear local result, however, is available. Start from a situation where $k=0$, i.e., actuarial pricing of insurance. Then it is well known that $a^* = L$, i.e., complete insurance (Mossin, 1968), so that $v(y;0) = u(y-pL)$. It can then be shown that

$$\frac{\partial A^v(y;k)}{\partial k}|_{k=0} = -\frac{\partial A^v(y;0)}{\partial y}pL,$$

that is, the increase in the loading has a mere income effect–the fact that a^* will be adjusted downwards is of second order. Under decreasing absolute risk aversion, the increase in loading thus increases the aversion towards the income risk.

C Imputation of risk aversion measures to the respondents

For each respondent in the sample, we want to estimate λ_i. To do so, we use information on both the chosen category $L_i \in \{[0, \frac{1}{2}], (\frac{1}{2}, \frac{2}{3}], (\frac{2}{3}, \frac{4}{5}], (\frac{4}{5}, 1]\}$ indicates the interval that respondent i has implicitly chosen for λ_i and the individual characteristics that we believe are related to risk taking.[37]

From (2), it follows that the density function for λ, conditional on the vector of characteristics x_i, is given by

$$g(\lambda|x_i, \widehat{\beta}, \widehat{\sigma}) = \frac{1}{\lambda(1-\lambda)} \frac{1}{\widehat{\sigma}} \phi\left(\frac{\log \frac{\lambda}{1-\lambda} - x_i'\widehat{\beta}}{\widehat{\sigma}}\right),$$

where $\phi(\cdot)$ is the pdf of the standard normal distribution and $(\widehat{\beta}, \widehat{\sigma})$ the ML estimates from Table 2 column (2).[38] Our estimate for the expected value of λ_i, given the characteristics and the respondent's stated choice of risk category L_i, is then

$$E(\lambda_i|x_i, \widehat{\beta}, \widehat{\sigma}, L_i) = \int_{\widetilde{\lambda} \in L_i} \widetilde{\lambda} \frac{g(\widetilde{\lambda}|x_i'\widehat{\beta}, \widehat{\sigma})}{\Pr(\lambda \in L_i|x_i'\widehat{\beta}, \widehat{\sigma})} \mathrm{d}\widetilde{\lambda}. \tag{7}$$

For each respondent, we compute (7) and denote it as $\widehat{\lambda}_i$. The descriptive statics for $\widehat{\lambda}_i$ are given in the upper panel of Table 3.

R_i and λ_i are related as in (3). This function is strictly increasing in R_i, but a closed form for the inverse does not exist. We approximate the inverse by $f(\lambda) \stackrel{\text{def}}{=} \frac{.2309\lambda + 2.6146\lambda^2 - 2.1585\lambda^3}{1-\lambda}$ where the coefficients in the numerator where estimated by NLLS using 100 data points generated by (2) for $R = 0.1, 0.2, \ldots, 10.0$ (all p-values equal 0.000, $R^2 = 1.0000$).

We then calculate $E(R_i|\cdot)$ as

$$E(R_i|\cdot) \simeq f(\widehat{\lambda}_i) + \frac{1}{2} f''(\widehat{\lambda}_i) \cdot var(\lambda_i|\cdot) + \frac{1}{6} f'''(\widehat{\lambda}_i) \cdot skew(\lambda_i|\cdot)$$

where the conditional mean is given by (7) and the conditional variance and skewness are calculated in a similar way. The results are given in the lower panel of Table 3.

[37] We follow Hsiao and Mountain (1985). The cost is that the estimate of $E(\lambda_i|x_i, L_i)$ has a variance that introduces noise in the regressor. To account for this noise we use bootstrapping methods.

[38] Thus we also control for educational attainment, civil status, labour market status, income scale, county of residence, type of residential area.

For the figures in column (1) of Table 3 we make use of the estimates for model (2) without covariates (given in column (1) of Table 2).

D Validation of the BKJS instrument: correlations between stated risk aversion and risk behaviours

Having established the cardinal measure for stated risk attitude, we now seek its validation by relating it to risk behaviours in two domains, health and occupation, and stated risk behaviour in the financial market.

We have estimated probit models for daily smoking (25.6 percent of respondents), being obese (BMI>30, 15.2 percent of respondents)[39], working in the private sector (55 percent of labour market active respondents), having a top manager position (11 percent of labour market active respondents), and an ordered probit model for the stated likelihood of stock investment with borrowed funds.[40] The results are presented in Table D.1.[41] In all five cases, our stated risk aversion measure has a significant negative correlation with the (stated) risk behaviour. Moreover these results are robust (see Tables H.2-H.5 in supplementary appendix H for the robustness of the estimates to sequential introduction of sets of controls). They are also in line with findings for other countries.[42] We should add, though, that the omission of the respondent's perception of risk as a control variable may bias the estimate of the coefficient with the risk aversion measure.[43,44]

[39] Respondents were asked about their height and weight. Mean BMI is 25.8 (st. dev. 4.4).

[40] The survey asked "How likely is it that you would borrow money to invest in stocks?" with answer alternatives: "very likely" (2 percent of the respondents), "likely" (3 percent), "somewhat likely" (21 percent), and "not likely at all" (74 percent).

[41] The standard errors reported in Table D.1 are obtained by bootstrapping to account for the fact that $\widehat{\lambda}$ is a generated regressor. From the original sample, 200 new samples with N observations are drawn with replacement. For each sample k, we estimate $(\widehat{\beta}^k, \widehat{\sigma}^k)$ using model (**??**) and the values for $E(\lambda_i|x_i'\widehat{\beta}^k, \widehat{\sigma}^k, L_i)$ using (7). Next, we estimate the risk behaviour model for each sample k. The reported standard errors are then the sample standard deviations of the 200 estimates for the last model.

[42] E.g., for the US BJKS report significant positive correlations between risk tolerance and smoking, drinking, being self-employed, having no health insurance. Refer also to the studies cited at the end of section 2. Bonin *et al.* (2007) find that the answer to the general risk tolerance question in the German SOEP survey has a significant positive correlation with the earnings risk measure (2-digit occupation level).

[43] If risk aversion and risk perception are the only variables affecting risk behaviour (discouraging it), then the sign of the bias is given by that of the covariance between risk aversion and perception.

[44] Few studies that measure the impact of risk aversion on risk behaviour control for risk perception. An exception is Lusk and Coble (2005) on the effect of an elicited risk aversion measure on the (stated) preferences for genetically modified (GM) food. These authors include a summary measure of three answers to risk perception questions. The risk aversion and risk perception measures are weakly correlated (.047) and the coefficient with the imputed CARA (based on answers to a multiple price list experiment) in the ordered probit regression for the

Table D1. Lifestyle choices, labour market choices and financial market choices

	Smoking[a]	Obese[a]	Private sector[a]	Top manager[a]	Loan[b]
$\widehat{\lambda}$	−.451* (.269)	−.533* (.295)	−.653** (.312)	−.959** (.344)	−.922*** (.252)
age	.067*** (.022)	.083*** (.026)	−.017*** (.004)	.017*** (.005)	−.030 (.025)
age^2	−.0007*** (.0002)	−.0009*** (.0003)	\	\	.0002 (.0003)
man	−.103 (.095)	.105 (.111)	.636*** (.088)	.558*** (.114)	.281*** (.088)
religious? yes[c]	−.243** (.101)	.251** (.107)	\	\	.149 (.100)
educational attainment	yes	yes	yes	yes	yes
civil status	yes	yes	yes	\	yes
income classes	yes	yes	\	\	yes
sector of work	yes	yes	\	\	yes
county of residence	yes	yes	yes	yes	yes
type residential area	yes	yes	yes	yes	yes
N	1486	1499	1080	995	1477
McKelvey & Zavoina's R^2	.186	.122	.285	.194	.169
Log lik	−763.44	−596.98	−632.84	−305.96	−1003.57

[a] Probit model, [b] Ordered probit model. [c] Dummies included for *religious? do not know* and *religious? do not want to answer*. Bootstrapped standard errors. Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test).

willingness to eat GM food falls in absolute value (from 8.26, s.e. 2.7 to 5.91, s.e. 2.24) when omitting the risk perception variable (personal communication with Jayson Lusk). Thus in that study the omission bias (probably due to correlation of risk perception with another variable, like gender) goes in the direction of the null hypothesis of no effect of risk aversion.

Conversely, few studies that are interested on the effect of risk perception on risk behaviour (e.g., Viscusi, 1990, on smoking; Kan and Tsai, 2004, on obesity) control directly for risk aversion.

7 Supplementary appendices

E Imputation of missing income category information

The survey asked respondents about their personal gross yearly income. Respondents could answer by ticking off one of 11 boxes. The answers are as shown in Table A2, columns 1-4. 79 respondents either did not want or could not provide income information (boxes 10 and 11, respectively). In addition, one respondent did not tick off any of the boxes.

In order not to miss 80 observations, we ran an OLS regression ($R^2 = .53$) of the integer in column 1 of Table A2 (I_i) on a set of socio-economic characteristics (age, age^2, gender, educational attainment, labour market status, sector of employment, type of position, self assessed health, area of residence, residential area type). We used the predicted I_i, $\widehat{I}_i$, for all respondents for which the *rhs* variables were available and imputed an income category, $\widetilde{I}_i$, defined as the nearest integer (i.e., $\widetilde{I}_i$ =nint($\widehat{I}_i$)) to each respondent. For the 1451 subjects who responded to the gross income question (i.e., for which $I_i \leq 9$) and for whom data on the *rhs* variables are available, the Spearman correlation between I_i and $\widetilde{I}_i$ is .7360, indicating a good fit.[45] For all subjects who responded to the gross income question, we kept their answer (I_i), while for the 80 subjects who did not respond but for which a prediction was possible (75 subjects) we imputed $\widetilde{I}_i$. The new income distribution is then as in columns 5-6 of Table A2.

Table E1. Response distribution to the question on gross yearly income

(1)	(2)	(3)	(4)	(5)	(6)
I_i	Gross yearly income in 1000 NOK	#	%	#	%
1	$\leq$ 100	127	8.18	148	9.55
2	100-199	185	11.91	197	12.75
3	200-299	346	22.28	364	23.50
4	300-399	368	23.70	379	24.47
5	400-499	199	12.81	209	13.49
6	500-599	108	6.95	110	7.10
7	600-799	84	5.41	85	5.49
8	800-999	35	2.25	35	2.26
9	>1000	22	1.42	22	1.42
10	Do not want to respond	51	3.28	0	0
11	Do not know	28	1.80	0	0
	Sum	1553	100	1549	100

[45] Estimation of an ordered probit model for income category resulted in a slightly worse correlation between I_i and $\widetilde{I}_i$ of .7182.

F Post-stratification of the average $\widehat{\lambda}$ and $\widehat{R}$

When we compare our sample with the 2006 population in terms of age, gender, educational attainment, and area of residence, it transpires that some groups are over/under-represented.[46] Figures F1-F3 show that the sample over-represents women in age group 30-34, men in age group 50-54, highly educated people, residents of Mid Norway, and under-represents women in age group 50-54, men in age group 60-64, people with low educational attainment, and residents of South-West Norway. Moreover, there is a slight over-representation of women in the sample (.511 in sample *vs* .497 in population).

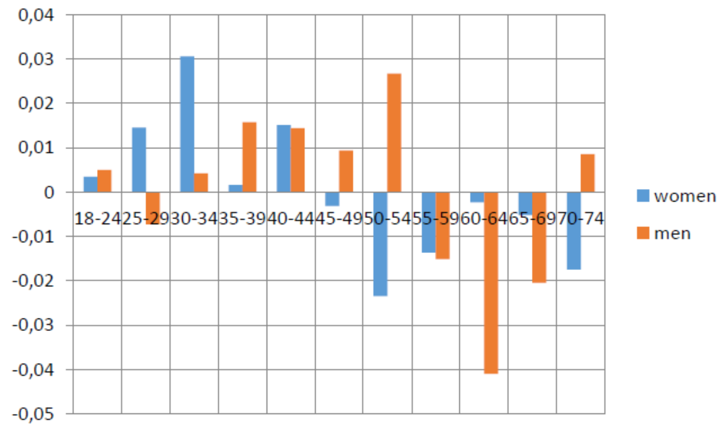

Figure F1. Over (+)- and under (−)-representation of women/men according to age group (sample fraction minus population fraction).

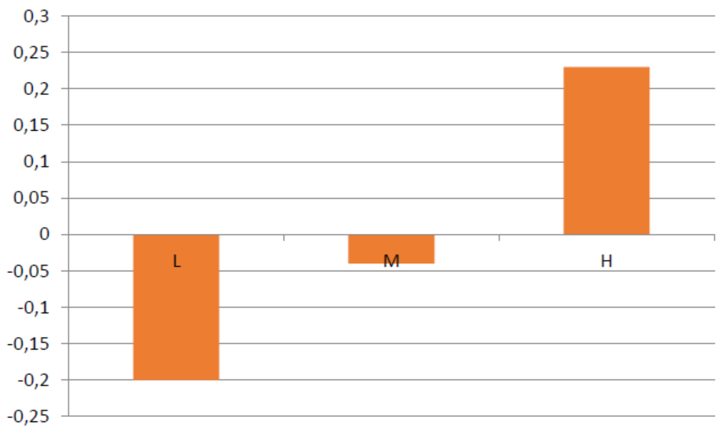

Figure F2. Over (+)- and under (−)-representation according to educational attainment (sample fraction minus population fraction).

[46] We are grateful to our colleague Aline Bütikofer for providing us with the 2006 population figures from register data.

Figure F3. Over (+)- and under (−)-representation of women/men according to region of residence (sample fraction minus population fraction).

Given the effects of gender and aging on risk aversion, we cannot rule out that the population means of $\widehat{\lambda}$ and $\widehat{R}$ are inconsistently estimated by means of the sample averages.

We have access to the 2006 population fractions for each of the 396 cells (or strata) spanned by the following categories: 2 genders, 3 levels of educational attainment, 11 age groups, 6 regions of residence.

With only 1509 observations for $\widehat{R}$, many of these cells are empty or nearly empty for the sample. Hence, post-stratification (or complete multi-way stratification) as a method to obtain a consistent estimate of the population average of $\widehat{R}$ is not feasible. Bethlehem and Keller (1987) proposed a linear weighting scheme with post-stratification as a special case. The scheme is feasible even with empty strata. It is an incomplete multi-way stratification method in the sense that does not make use of all information contained in the population distribution of observable characteristics, but rather employs only the marginal distribution for some characteristics. For our purpose, we have chosen the bivariate distribution of age and gender (22 cells), and the marginal distributions of educational attainment and of region of residence (3 and 6 cells, respectively).

The weighting scheme consists of regressing $\widehat{R_i}$ ($i = 1, \ldots, 1509$) on a constant and 22+3+6 dummy variables (after dropping one dummy variable for each distribution to avoid multi-collinearity). Next, the obtained vector of coefficients is multiplied with the vector of corresponding population fractions (augmented with unity) to obtain a consistent estimate of the population average of $\widehat{R}$. It can easily be shown that in case the complete distribution were used (in our case using $22 \times 3 \times 6$ dummies), the method is equivalent to standard post-stratification.

When we use the Bethlehem-Keller linear weighting scheme, we obtain estimates for the population means of $\widehat{\lambda}$ and $\widehat{R}$ that are very close to the sample averages: .733 and 3.831, respectively.

G Robustness check for model (2) and the risk behaviour models

Table G1 shows the results for the interval regression estimates for model (2) when sequentially introducing control variables. Tables G2-G8 show the maximum likelihood estimates for the risk behaviour models when sequentially introducing control variables. The standard errors reported in Tables G2-G8 are standard ones, i.e., not obtained through bootstrapping.

Table G1 Robustness analysis interval regression (**??**)

variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
age	.002	.002	.002	.000	.001	.000	.001	.000	.001
	(.002)	(.002)	(.002)	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)
man	−.639***	−.643***	−.642***	−.650***	−.684***	−.705***	−.698***	−.703***	−.695***
	(.158)	(.158)	(.158)	(.160)	(.160)	(.161)	(.160)	(.160)	(.160)
man×age	.009***	.010***	.009***	.010***	.010***	.011***	.011***	.011***	.011***
	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)
educational attainment	no	yes	yes	yes	yes	yes	yes	yes	yes
income classes	no	no	yes	yes	yes	yes	yes	yes	yes
civil status	no	no	no	yes	yes	yes	yes	yes	yes
labour mkt status	no	no	no	no	yes	yes	yes	yes	yes
county	no	no	no	no	no	yes	yes	yes	yes
type residential area	no	no	no	no	no	no	yes	yes	yes
religiosity	no	no	no	no	no	no	no	yes	yes
life satisfaction	no	no	no	no	no	no	no	no	yes
constant	1.151***	1.193***	1.223***	1.352***	1.430***	1.590***	1.352***	1.331***	1.223***
	(.104)	(.137)	(.154)	(.172)	(.195)	(.221)	(.242)	(.243)	(.249)
σ	.862***	.861***	.860***	.859***	.858***	.851***	.848***	.846***	.845***
	(.029)	(.029)	(.029)	(.029)	(.029)	(.029)	(.029)	(.029)	(.029)
N	1509	1509	1509	1509	1509	1509	1509	1509	1509
McKelvey & Zavoina's R^2	0.019	0.021	0.022	0.023	0.024	0.026	0.027	0.030	0.030
Log lik	−1911.77	−1911.22	−1909.32	−1907.80	−1905.72	−1896.49	−1892.43	−1889.14	−1886.32

Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test).

Table G2. Robustness analysis probit model for smoking

variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(9)
$\widehat{\lambda}$	$-.422^{*}$ (.230)	$-.409^{*}$ (.231)	$-.529^{**}$ (.237)	$-.509^{**}$ (.238)	$-.504^{**}$ (.240)	$-.501^{**}$ (.241)	$-.480^{**}$ (.245)	$-.499^{**}$ (.246)	$-.451^{*}$ (.247)	b
age	$.058^{***}$ (.015)	$.059^{***}$ (.015)	$.067^{***}$ (.016)	$.090^{***}$ (.018)	$.078^{***}$ (.020)	$.071^{***}$ (.021)	$.068^{***}$ (.021)	$.069^{***}$ (.021)	$.067^{***}$ (.021)	$.067^{***}$ (.021)
age^2	$-.0006^{***}$ (.0002)	$-.0006^{***}$ (.0002)	$-.0008^{***}$ (.0002)	$-.0010^{***}$ (.0002)	$-.0008^{***}$ (.0002)	$-.0008^{***}$ (.0002)	$-.0007^{***}$ (.0002)	$-.0007^{***}$ (.0002)	$-.0007^{***}$ (.0002)	$-.0007^{***}$ (.0002)
man	$-.141^{**}$ (.071)	$-.171^{**}$ (.072)	$-.165^{**}$ (.073)	$-.143^{*}$ (.075)	$-.137^{*}$ (.080)	$-.112$ (.083)	$-.107$ (.084)	$-.104$ (.084)	$-.103$ (.084)	$-.097$ (.084)
religious? yes^a	\	$-.324^{***}$ (.087)	$-.284^{***}$ (.090)	$-.249^{***}$ (.091)	$-.255^{***}$ (.091)	$-.253^{***}$ (.093)	$-.250^{***}$ (.094)	$-.247^{***}$ (.094)	$-.244^{***}$ (.094)	$-.249^{***}$ (.094)
educational attainment	no	no	yes	yes	yes	yes	yes	yes	yes	yes
civil status	no	no	no	yes	yes	yes	yes	yes	yes	yes
income classes	no	no	no	no	yes	yes	yes	yes	yes	yes
sector of work	no	no	no	no	no	yes	yes	yes	yes	yes
county	no	no	no	no	no	no	yes	yes	yes	yes
type residential area	no	no	no	no	no	no	no	yes	yes	yes
life satisfaction	no	no	no	no	no	no	no	no	yes	yes
constant	-1.447^{***} (.352)	-1.403^{***} (.354)	$-.982^{***}$ (.379)	-1.768^{***} (.442)	-1.788^{***} (.458)	-1.674^{***} (.464)	-1.597^{***} (.493)	-1.692^{***} (.525)	-1.520^{***} (.535)	-1.746^{***} (.516)
N	1492	1492	1492	1492	1492	1486	1486	1486	1486	1486
Pseudo R^2	0.0121	0.0211	0.0583	0.0712	0.0754	0.0833	0.0903	0.0916	0.0964	.0970
Log lik	-837.41	-829.78	-798.26	-787.28	-783.78	-774.47	-768.53	-767.42	-763.44	-762.93

a dummies included for *religious? do not know* and *religious? do not want to answer*. b: model (9) uses dummies for risk aversion categories (reference category is group 1 ("Job 2","Job 2")), instead of $\widehat{\lambda}$. The coefficients are as follows: group 2: $-.053(.163)$, group 3: $-.069$ (.122), group 4: $-.209^{*}$ (.121). Wald test for all three coefficients zero: $\chi^2(3) = 4.37, p = .2241$. Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test).

Table G3. Robustness analysis probit model for obesity

variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(9)
$\widehat{\lambda}$	−.452*	−.465*	−.493*	−.502*	−.522**	−.513*	−.539**	−.538**	−.533**	b
	(.263)	(.265)	(.267)	(.267)	(.265)	(.267)	(.270)	(.270)	(.272)	
age	.059***	.059***	.064***	.065***	.070***	.081***	.084***	.083***	.083***	.082***
	(.018)	(.018)	(.019)	(.021)	(.024)	(.025)	(.025)	(.025)	(.025)	(.025)
age^2	−.0006***	−.0006***	−.0006***	−.0007***	−.0007***	−.0009***	−.0009***	−.0009***	−.0009***	−.0009***
	(.0002)	(.0002)	(.0002)	(.0002)	(.0003)	(.0003)	(.0003)	(.0003)	(.0003)	(.0003)
man	.034	.058	.071	.063	.097*	.107	.107	.104	.105	.113
	(.081)	(.081)	(.082)	(.083)	(.089)	(.090)	(.091)	(.091)	(.091)	(.091)
religious? yes[a]	\	.252***	.271***	.250***	.249***	.258***	.250***	.251***	.251***	.249***
		(.093)	(.094)	(.095)	(.096)	(.096)	(.096)	(.096)	(.096)	(.096)
educational attainment	no	no	yes	yes	yes	yes	yes	yes	yes	yes
civil status	no	no	no	yes	yes	yes	yes	yes	yes	yes
income classes	no	no	no	no	yes	yes	yes	yes	yes	yes
sector of work	no	no	no	no	no	yes	yes	yes	yes	yes
county	no	no	no	no	no	no	yes	yes	yes	yes
type residential area	no	no	no	no	no	no	no	yes	yes	yes
life satisfaction	no	no	no	no	no	no	no	no	yes	yes
constant	−2.068***	−2.120***	−2.213***	−2.116***	−2.207***	−2.297***	−2.235***	−2.165***	−2.125***	−2.335***
	(.416)	(.420)	(.455)	(.515)	(.538)	(.551)	(.584)	(.604)	(.611)	(.596)
N	1505	1505	1505	1505	1505	1499	1499	1499	1499	1499
Pseudo R^2	0.0137	0.0203	0.0298	0.0365	0.0411	0.0525	0.0546	0.0631	0.0633	.0634
Log lik	−629.60	−625.40	−619.30	−615.04	−612.08	−603.91	−602.53	−597.16	−596.98	−596.94

a dummies included for *religious? do not know* and *religious? do not want to answer*. *b*: model (9) uses dummies for risk aversion categories (reference category is group 1 ("Job 2","Job 2")), instead of $\widehat{\lambda}$. The coefficients are as follows: group 2: −.149 (.178), group 3: −.163 (.130), group 4: −.255* (.131). Wald test for all three coefficients zero: $\chi^2(3) = 3.89, p = .273$. Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test).

Table G4. Robustness analysis probit model for private sector employment

variable	(1)	(2)	(3)	(4)	(5)	(6)
$\widehat{\lambda}$	$-.407$	$-.604^{**}$	$-.598^{**}$	$-.655^{**}$	$-.652^{**}$	a
	(.259)	(.268)	(.269)	(.274)	(.275)	
age	$-.014^{***}$	$-.015^{***}$	$-.016^{***}$	$-.017^{***}$	$-.017^{***}$	$-.016^{***}$
	(.003)	(.003)	(.004)	(.004)	(.004)	(.004)
man	$.566^{***}$	$.607^{***}$	$.610^{***}$	$.635^{***}$	$.636^{***}$	$.628^{***}$
	(.079)	(.083)	(.084)	(.084)	(.084)	(.085)
educational attainment	no	yes	yes	yes	yes	yes
civil status	no	no	yes	yes	yes	yes
county	no	no	no	yes	yes	yes
type residential area	no	no	no	no	yes	yes
constant	$.780^{***}$	1.435^{***}	1.523^{***}	1.626^{***}	1.728^{***}	1.608^{***}
	(.226)	(.301)	(.331)	(.365)	(.399)	(.363)
N	1080	1080	1080	1080	1080	1083
McKelvey & Zavoina's R^2	0.104	0.229	0.234	0.271	0.278	0.285
Log lik	-702.32	-652.67	-650.47	-637.14	-634.37	-632.84

a: model (6) uses dummies for risk aversion categories (reference category is group 1 ("Job 2","Job 2")), instead of $\widehat{\lambda}$. The coefficients are as follows: group 2: $-.285$ (.182), group 3: $-.504^{***}$ (.136), group 4: $-.344^{**}$ (.136). Wald test for all three coefficients zero: $\chi^2(3) = 14.11, p = .0028$. Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test).

Table G5. Robustness analysis probit model for top manager

variable	(1)	(2)	(3)	(4)	(5)
$\widehat{\lambda}$	-1.042^{***}	-1.012^{***}	-0.976^{***}	-0.956^{***}	a
	(.349)	(.353)	(.366)	(.367)	
age	$.016^{***}$	$.016^{***}$	$.017^{***}$	$.017^{***}$	$.017^{***}$
	(.005)	(.005)	(.005)	(.005)	(.005)
man	$.520^{***}$	$.542^{***}$	$.554^{***}$	$.558^{***}$	$.535^{***}$
	(.114)	(.115)	(.118)	(.119)	(.120)
educational attainment	no	yes	yes	yes	yes
county	no	no	yes	yes	yes
residential area type	no	no	no	yes	yes
constant	-1.489^{***}	-2.017^{***}	-2.326^{***}	-2.118^{***}	-2.597^{***}
	(.304)	(.409)	(.482)	(.522)	(.475)
N	1018	1002	995	995	998
McKelvey & Zavoina's R^2	.122	.137	.186	.194	.203
Log lik	-323.52	-318.48	-307.16	-305.96	-303.37

a: model (5) uses dummies for risk aversion categories (reference category is group 1 ("Job 2","Job 2")), instead of $\widehat{\lambda}$. The coefficients are as follows: group 2: .263 (.216), group 3: $-.273$ (.179), group 4: $-.336^{*}$ (.181). Wald test for all three coefficients zero: $\chi^2(3) = 12.45, p = .006$. Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test).

Table G6. Robustness analysis ordered probit model for willingness to take up loan.

variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
$\widehat{\lambda}$	−.804*** .221	−.832*** (.222)	−.829*** (.222)	−.831*** (.223)	−.838*** (.221)	−.845*** (.221)	−.811 (.226)	−.850*** (.229)	−.922*** (.230)	b
age	−.017 .015	.016 (.015)	.007 (.016)	−.006 (.017)	−.046** (.019)	−.042** (.022)	−.034 (.022)	−.030 (.022)	−.030 (.020)	−.030 (.023)
age^2	−.000 .000	−.000 (.000)	−.000 (.000)	−.000 (.000)	−.000 (.000)	−.000 (.000)	−.000 (.000)	.000 (.000)	.000 (.000)	.000 (.000)
man	.388*** .070	.402*** (.070)	.408*** (.070)	.405*** (.071)	.271*** (.074)	.265*** (.076)	.272*** (.079)	.262*** (.079)	.267*** (.079)	.249*** (.079)
religious? yes^a	\	.168** (.082)	.152* (.083)	.143* (.083)	.138 (.084)	.139* (.084)	.142* (.086)	.154* (.086)	.149* (.086)	.154* (.087)
educational attainment	no	no	yes	yes	yes	yes	yes	yes	yes	yes
civil status	no	no	no	yes	yes	yes	yes	yes	yes	yes
income classes	no	no	no	no	yes	yes	yes	yes	yes	yes
labour mkt status	no	no	no	no	no	yes	yes	yes	yes	yes
sector of work	no	no	no	no	no	no	yes	yes	yes	yes
county	no	no	no	no	no	no	no	yes	yes	yes
residential area type	no	no	no	no	no	no	no	no	no	yes
cut1	.536 .341	.550 (.345)	.660 (.380)	.252 (.448)	−.249 (.462)	−.319 (.529)	−.012 (.568)	.191 (.596)	.602 (.638)	1.013 (.622)
cut2	1.565 .341	1.580 (.345)	1.698 (.380)	1.294 (.448)	.818 (.462)	0.751 (.530)	1.103 (.569)	1.289 (.598)	1.703 (.640)	2.123 (.626)
cut3	2.020 .343	2.039 (.346)	2.159 (.381)	1.760 (.447)	1.291 (.461)	1.227 (.529)	1.573 (.570)	1.763 (.598)	2.175 (.641)	2.595 (.625)
N	1483	1483	1483	1483	1483	1483	1477	1477	1477	1480
McKelvey & Zavoina's R^2	.058	.063	.075	.082	.116	.122	.142	.157	.169	.175
Log lik	-1062.95	-1060.72	-1055.84	-1052.44	-1033.70	-1031.23	-1015.14	-1008.34	-1003.57	-1000.05

a dummies included for *religious? do not know* and *religious? do not want to answer*. b: model (9) uses dummies for risk aversion categories (reference category is group 1 ("Job 2","Job 2")), instead of $\widehat{\lambda}$. The coefficients are as follows: group 2: .164 (.134), group 3: $-.354^{***}$ (.114), group 4: $-.356^{***}$ (.111). Wald test for all three coefficients zero: $\chi^2(3) = 28.36, p = .0000$. Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test).

H Sources of and comments on constructing the figures in Table 6

(1.1-3) and (2.1-3) CWED described in Scruggs *et al.* (2014a). The figures for Chile are computed using the CWED code (Scruggs *et al.* (2014b) on the Chilean numbers taken from Social Security Administration (2003). The coverage numbers are proxied by the coverage rate for social security for Chile in 2003 listed Rofman (2005: Table 1). For Italy, CWED applies the UB coverage number for 1985 (46%, attributed in Scruggs *et al.* (2014b: 56) to Flora (ed.) 1986, *Growth to Limits*, Vol 4, p. 522) to all subsequent years. We therefore replace this coverage rate by the 2008 number available from Banca d'Italia (2009: 102-3), i.e., 88%, which changes the unemployment generosity index from 5.8 to $5.8 \times .88/.46 = 11.1$.

(1.4) Vroman's (2001, 2003) UB generosity index computed as $\frac{\text{total unempl comp}}{\text{wage bill}} \times \frac{1-p_u}{p_u}$. Total unemployment compensation is taken from OECD.Stat (Social protection and well-being: social expenditure. Branch: unemployment. Type of programme: unemployment: unemployment compensation / severance pay) (SOCX database). Wage bill is taken from OECD.Stat (Annual national accounts) and p_u is the harmonised unemployment rate (see (1.5)). For FR and NL, the amounts include both unemployment insurance (UI) and unemployment assistance (UA) benefits. For NL the amount includes all assistance benefit spending, including payments to inactive recipients (see OECD (2014) "The scope and comparability of data on labour market programmes" available at www.oecd.org/els/emp/ALMPdata-Scope-and-Comparability-2015.pdf). For this reason, the amounts for FR and NL are adjusted downwards by multiplying them with the ratio of UI benefit recipients to UI+UA recipients (taken from OECD.Stat (Labour: Labour market programmes)).

(1.5) OECD.Stat (Labour force statistics).

(2.4-5) Measured in full-time equivalent paid sick days. Source: Heymann *et al.* (2009) supplemented with numbers for Chile taken from Social Security Administration (2003).

(2.6) The Vroman-like sickness generosity index computed as $\frac{\text{paid sickness leave comp}}{\text{wage bill}} \times \frac{1-p_s}{p_s}$. Paid sickness leave compensation is taken from OECD.Stat (Social protection and well-being: social expenditure. Branch: Incapacity related. Type of programme: incapacity related – paid sick leave (other sickness daily allowances)) (SOCX database). Wage bill is taken from OECD.Stat (Annual national accounts) and p_s is the fraction of statutory working days absent due to illness (see (2.7)).

(2.7) OECD.Stat (Health: health status: absence from work due to illness). For CL and US: self-reported; for CL: 2003 since 2002 number is not available. For NL, FR, IT and NO: *compensated* absence. These numbers are subsequently augmented with the "deductible" for a 5-day flu episode (as given by (2.4)) in order to obtain the total number of days absent from work due to illness.

(2.8) Ratio of (2.7) to the number of days at work. The latter is calculated as

261 days ($365 - 2 \times 52$) *minus* the number of days minimum annual leave (listed at https://en.wikipedia.org/wiki/List_of_minimum_annual_leave_by_country) *minus* (2.7).

(3.1) Ratio of "expenditure on curative and rehabilitative care" financed under "government schemes and compulsory contributory financing schemes" to the total expenditure on the same care, OECD.Stat (Health: Health expenditure and financing); for CL: percentage of total health spending by public sector, taken from Bitrán and Urcullo (2008: 102).

(3.2) Fraction of population covered under government/social health insurance OECD.Stat (Health: Social protection)

(3.3) Complement of the fraction of the population covered under public and primary private health insurance (OECD.Stat (Health: Social protection: Total public and primary private health insurance).

(3.4)-(3.5) Five-year interval rates (1×5) for both sexes. CL, US, FR and NL: 2000-04; IT and NO: 2005-09. Source: Human Mortality Database.

I Short descriptions of the six health insurance systems

In **the Netherlands** in 2004, all employees and self-employed with earnings below a threshold are obliged to participate in the statutory health insurance (SHI) system consisting of competing sickness funds. They pay a regulated income-related premium and a flat rate contribution determined by the sickness fund; this also covers their dependents. Benefits are in-kind. In 2004, SHI covered 65% of the population. Citizens with higher earnings can seek health insurance coverage in the private health insurance system (PHI). The PHI covered about 26% of the population. There is free contracting about the premium and the health services that are reimbursed, and both can be contingent on the insuree's medical history. Co-payments and deductibles apply. Coverage of dependents requires the payment of extra premia. In addition to the risk-rated premium, the insuree pays solidarity contributions to cross-subsidise the SHI. In addition to the general PHI system, there exists two extra branches: (i) the WTZ branch (4.3% of the population) which offers a reimbursement insurance for a basic package of benefits that people who are not covered under SHI can buy if their premium under the PHI is above a critical level (not for family members), (ii) the KPZ branch which covers civil servants (5%). (Source: Ministry of Health, Welfare and Sport, 2004.). In 2003, the mean waiting time in Dutch hospitals was 5.5 weeks for in-patient treatment and 5.1 weeks for out-patient treatment (Source: Schut and Varkevisser, 2013)

Chile in 2002 had a dual health system: a mix of a public-integrated model and a private insurance/provider model. The SHI (*Fonasa*) covered 67.1% of the population in 2002, while private insurers (*Isapres*) covered around 17.1% of the population.[47] Dependent workers are required by law to seek HI under either

[47] In 2005, 3.5% is covered under a scheme for the armed forces or universities, while 10.5%

SHI or PHI. The Fonasa premium was fixed at 7% of income, except for indigent people who were covered by Fonasa at a zero premium. Fonasa provides health care through public providers;[48] it does not provide prescription drugs insurance. Waiting lists are common for public specialist care (OECD 2003: 115). The Isapres premium is legally bounded from below by 7% income premium charged by Fonasa. The Isapres plans are individually contracted with prices, benefits and co-payment rates (in the range of 30 to 50%) determined by applicant's health risk, age and sex. The PHI companies do not receive subsidies from the government and the lack of *ex ante* risk-adjustment schemes has led to extreme adverse selection (OECD 2003:115). Hence, PHI companies engage in cream-skimming and risk-selection to guarantee profitability (Bitrán and Urcullo, 2008, and Holst *et al.* 2004). This results in a PHI membership profile of people belonging to the higher income deciles and below the age of 50. In a large survey ($N = 33529$) carried out in 2000, 30% of the respondents answered that they had unmet health needs (30% of Fonasa members, 23.6% of Isapres members) (see Frenz *et al.* 2014).

France has a universal public health insurance covering almost 100% of the population. It covers hospital care, ambulatory care and prescription drugs. Citizens have substantial freedom of choice in selecting health care providers. Waiting lists are of minor importance. Co-insurance rates apply to all health services and drugs listed in the publicly financed benefit package. Also co-payments apply (up to an annual ceiling). Exemptions from co-insurance apply (chronically ill, people with income below a threshold, people receiving invalidity and work-injury benefits). The incompleteness of the cover by the public insurance system has triggered a market for supplementary PHI which reimburses statutory cost-sharing. The market for PHI is made up of *mutuelles* (60% market share), provident institutions (19%) and private insurance companies (21%). The mutuelles and provident institutions operate on a solidarity principle in that their premia are community rated and not based on the risk of the individual applicants. In 2002, 87.5% of the French population had supplementary PHI. (Source: Buchmueller and Coufinhal, 2004; Commonwealth Fund, 2010.)

In the **US**, public health insurance is available to retired citizens and some of the disabled under age 65 (Medicare) and to citizens with low incomes (Medicaid). The basic Medicare plan (part A) provides hospital insurance. In 2002, this could be supplemented with Part B (covering additional health care services) (Part D, that covers outpatient prescription drugs, was not introduced before 2006). Medicare Advantage plans (sometimes called Medicare Part C) are offered by private insurers for those retirees that wish to opt out of the traditional Medicare

remained uninsured (Bitrán and Urcullo, 2008:106).

[48]Except for members who have pruchased a "free choice" policy–they can directly contact private providers.

plans. Additional supplementary plans (Medigap and employment-based health benefit plans) are available for extending the coverage of Parts A and B. For people under the age of 65, 83% had in 2002 HI coverage (64.2% employment-based HI, 15.9% public-sector HI, 6.7% individually purchased HI) while 17% had no HI coverage. For workers the numbers are 72% with coverage (72.4%, 6.3%,5.8%) while 18% was uninsured, 60% of which where self-employed or working in private sector firms with less than 100 employees. Simulations run by the Employee Benefit Institute indicate that in 2002, workers with an employment-based health benefit plan extending to retirement will need to save in the range $37,000-$75,0000 by the age of 65 to pay for the health insurance premia and out-of-pocket payments in retirement during the rest of life (the amounts depend on choice of supplementary Medicare plan, age of death, and interest rate), while those without such a benefit plan would need to save in the range $47,000-$1,458,000 by the age of retirement (Fronstin and Salisburry, 2003).

In **Norway** 2006, the entire population was covered by the public health insurance system, which is tax financed. Under the system, patients can freely choose provider. Specialised health care (including pharmaceuticals) in public hospitals is free of charge. Co-payments apply for primary health care and for medicine but their sum was capped at 1,615NOK in 2006 (about 182USD or 150€, PPP). There are significant waiting lists for elective treatment: average waiting time for specialist health care (all procedures) was 72 days in 2007. In order to insure oneself against long waiting times, a small fraction (<1%) of the population in 2006 took out a PHI that gave them the right to be treated in a private hospital (either in Norway or abroad) in case no public treatment is available within 2 or 3 weeks after diagnosis. (Source: Commonwealth Fund, 2010; Askildsen, Iversen and Kaarbøe, 2013.)

Italy, like Norway, has a national health service that is mainly tax-financed. In-patient care and GP services are free of charge. Co-payments are required on pharmaceuticals, diagnostic procedures and specialist visits. No general co-payment cap exists, but there are many exemptions (e.g., for people with a family income below a threshold, for chronically ill people). Waiting times prevail, but those for in-patient care appear less critical than those for out-patient care and diagnostics. The average waiting time for diagnostic tests was 50 days in 2009. Around 15% of the population had in 2006 some form of private health insurance for covering co-payments, for direct and faster access to specialists and diagnostic services, for extended choice of hospitals and clinics and access to dental care. (Source: Commonwealth Fund, 2010; Fattore *et al.*, 2013).

J Exploring the consequences of strong endogeneity of life satisfaction and religiosity

Variables like life satisfaction and religiosity are often included as covariates in risk attitude equations (see, e.g., Dohmen *et al.* (2005, 2011) but their endogeneous nature can bias the coefficients. Both variables are not easy to instrument for. Sometimes, climate data (hours of sunshine, cloud cover rate) have been used (Barrington-Leigh, 2008, Guven 2012, Guven and Hoxha, 2015). Unfortunately, we could not make use of such a climate instrument because (i) the survey was done in a short time span and (ii) we only have information in which county each respondent lives but not in which city/town/village. Since Norwegian counties are geographically stretched and the topography leads to a lot of intracounty climate variation, just taking an average of measured sunshine hours or cloud cover for each county is bound to be a weak instrument due to lack of variation. We have no instrument for religiosity either–a variable that is notoriously difficult to instrument for (see, e.g., Hungerman, 2014: 1054).

Having no instrument available for either life satisfaction and religiosity, we have asked the question raised by Altonji *et al.* (2005) "suppose that selection on unobservables (SoU) is as large as selection on observables (SoO), what is then the coefficient with the treatment variable?" (in our case, religiosity and life satisfaction). The Altonji *et al.* approach consists in replacing the exclusion restriction by a restriction on the correlation between the residuals in a bivariate model. Thus we ran two sets of bivariate models of the type

$$\begin{aligned} z^* &= x'\beta + u, \\ y^* &= \alpha z + x'\gamma + \varepsilon, \end{aligned}$$

with $\binom{u}{\varepsilon} \sim N(\binom{0}{0}, \begin{pmatrix} 1 & \rho\sigma \\ \rho\sigma & \sigma^2 \end{pmatrix})$. Here z^* is the latent variable for religiosity/life satisfaction, z is its observable counterpart and y^* is the latent variable for risk aversion. We estimated these models under the restriction that $\frac{cov(x'\beta,x'\gamma)}{var(x'\gamma)} = \frac{cov(z^*,\varepsilon)}{var(\varepsilon)}$, which corresponds to "SoU is as large as SoO".[49] Convergence was obtained only for a limited x-vector.

For life satisfaction, we used a three valued observable variable for z^* (0 if low or medium life satisfaction, 1 if high, 2 if very high) and an interval coded variable for y^* (as in the main text). The bivariate model (ordered probit for z^*, interval regression for y^*) was estimated using the `cmp` Stata module written by David Roodman (2011). Table J1 gives the results for the y^* equation when ρ is restricted to zero, while Table J2 gives the results when imposing the constraint that $\rho = \frac{cov(x'\beta,x'\gamma)}{var(x'\gamma)}$. Under the latter assumption, $\rho = .197$ and the coefficients with the dummies for high and very high life satisfaction change from positive/zero to negative. The coefficients with the included x-variables (age and gender, education, region, type of residential area) are hardly affected.

[49]Thanks to Todd Elder for providing us with the Stata code.

Table J1. Interval regression for λ assuming $\rho = 0$.

male	age	male×age	edu1	edu2	edu3	edu4
−.655**	.003	.010***	−.002	−.070	−.060	−.094
(.149)	(.002)	(.003)	(.085)	(.084)	(.090)	(.132)
East	South-West	West	Mid	North	ra1	ra2
.017	.041	−.067	−.062	−.047	.259**	.336***
(.107)	(.119)	(.103)	(.105)	(.129)	(.107)	(.115)
ra3	ra4	ls1	ls2	constant	σ	ρ
.216*	.330***	.119**	−.006	.850***	.856***	0
(.112)	(.119)	(.058)	(.076)	(.206)	(.026)	\

edu1: between 11 and 13 years; edu2: university short; edu3:university long; edu4: currently studying (reference: ≤ 10 years of education); reference for region: Oslo; ra1: residential area=country side; ra2: small town; ra3: medium town; ra4: big town (reference: farm); ls1: high life satisfaction; ls2: very high life satisfaction (reference: low or medium life satisfaction). Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test). $N = 1509$. Log lik: -3337.969.

Table J2. Interval regression for λ assuming $\rho = \frac{cov(x'\beta, x'\gamma)}{var(x'\gamma)}$.

male	age	male×age	edu1	edu2	edu3	edu4
−.631**	.004	.009***	−.024	−.084	.057	−.082
(.151)	(.003)	(.003)	(.086)	(.085)	(.092)	(.134)
East	South-West	West	Mid	North	ra1	ra2
.000	.050	−.062	−.060	−.037	.266**	.330***
(.108)	(.121)	(.104)	(.107)	(.131)	(.108)	(.116)
ra3	ra4	ls1	ls2	constant	σ	ρ
.220*	.339***	−.120**	−.483***	1.033***	.869	.197
(.113)	(.121)	(.058)	(.078)	(.208)	(.026)	\

edu1: between 11 and 13 years; edu2: university short; edu3:university long; edu4: currently studying (reference: ≤ 10 years of education); reference for region: Oslo; ra1: residential area=country side; ra2: small town; ra3: medium town; ra4: big town (reference: farm); ls1: high life satisfaction; ls2: very high life satisfaction (reference: low or medium life satisfaction). Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test). $N = 1509$. Log lik: -3337.9635.

For religiosity, we used a binary observable variable z (1 if "yes", 0 if "no" or "don't know") and a binary observable variable for y^* (1 if belonging to the two highest risk aversion groups, 0 otherwise); hence σ was normalised to 1.[50] The bivariate model was estimated by `biprobit`. Table H3 gives the results for the y^* equation when ρ is restricted to zero, while Table J4 gives the results when imposing the constraint that $\rho = \frac{cov(x'\beta, x'\gamma)}{var(x'\gamma)}$. Under the latter assumption, $\rho = .481$ and the coefficient with the religiosity dummy changes from positive to negative. The coefficients with the included x-variables (age and gender, education, region, type of residential area) are affected in a minor way.

[50] Convergence was not achieved when y^* was intervalcoded.

Table J3. Probit model for $1(\lambda > \frac{2}{3})$ assuming $\rho = 0$.

male	age	male×age	edu1	edu2	edu3	edu4
$-.743^{***}$	.019	.006	$-.318^{**}$	$-.414^{***}$	$-.496^{***}$	$-.410^{**}$
(.231)	(.004)	(.005)	(.155)	(.154)	(.161)	(.208)
East	**South-West**	**West**	**Mid**	**North**	**ra1**	**ra2**
.110	.055	$-.031$	$-.018$	.108	.234	$.306^{*}$
(.164)	(.185)	(.158)	(.163)	(.202)	(.164)	(.170)
ra3	**ra4**	**religious**	**constant**		σ	ρ
.110	$.420^{**}$	$.150^{*}$	.257		1	0
(.170)	(.181)	(.090)	(.322)		\	\

edu1: between 11 and 13 years; edu2: university short; edu3:university long; edu4: currently studying (reference: ≤ 10 years of education); reference for region: Oslo; ra1: residential area=country side; ra2: small town; ra3: medium town; ra4: big town (reference: farm). Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test). $N = 1509$. Log lik: -1562.993.

Table J4. Probit model for $1(\lambda > \frac{2}{3})$ assuming $\rho = \frac{cov(x'\beta, x'\gamma)}{var(x'\gamma)}$.

male	age	male×age	edu1	edu2	edu3	edu4
$-.593^{***}$	$.023^{***}$	.002	$-.259^{*}$	$-.304^{**}$	$-.391^{**}$	$-.379^{*}$
(.223)	(.004)	(.005)	(.150)	(.149)	(.155)	(.202)
East	**South-West**	**West**	**Mid**	**North**	**ra1**	**ra2**
.069	.124	$-.039$	$-.057$	.109	.246	$.322^{*}$
(.159)	(.178)	(.153)	(.158)	(.195)	(.159)	(.171)
ra3	**ra4**	**religious**	**constant**		σ	ρ
.144	$.390^{**}$	$-.699^{***}$	.209		1	.481
(.165)	(.176)	(.083)	(.312)		\	\

edu1: between 11 and 13 years; edu2: university short; edu3:university long; edu4: currently studying (reference: ≤ 10 years of education); reference for region: Oslo; ra1: residential area=country side; ra2: small town; ra3: medium town; ra4: big town (reference: farm). Statistical significance at the 1/5/10 percent level is denoted with ***/**/* (two-sided t test). $N = 1509$. Log lik: -1563.0137.

Thus if selection on unobservables is as large as on observables, the coefficients with variables like life satisfaction and religiosity are biased in the univariate model for risk aversion. A prudent guess is that some selection on ubservables is present and that therefore the coefficients with these variables are closer to zero. In Table J5 we replicate Table 3 of the main text, but now based on ML estimates for model (2) without life satisfaction and religiosity as controls. The results are very close to those in Table 3.

Table J5. Conditional sample distributions of $\widehat{\lambda}_i$ and $\widehat{R}_i$ w/o religiosity and life satisfaction as controls

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	mean^a	mean^b	median^b	st.dev.^{b}	min^b	max^b	N
$\widehat{\lambda}_i\|_{L_i=(0,\frac{1}{2})}$	.401	.401	.402	.011	.371	.417	201
$\widehat{\lambda}_i\|_{L_i=(\frac{1}{2},\frac{2}{3})}$	.595	.594	.594	.003	.587	.600	130
$\widehat{\lambda}_i\|_{L_i=(\frac{2}{3},\frac{4}{5})}$	.738	.738	.739	.002	.731	.742	621
$\widehat{\lambda}_i\|_{L_i=(\frac{4}{5},1)}$	.873	.872	.873	.005	.855	.882	557
$\widehat{\lambda}_i$	.731^c	.730^c	.740	.154	.371	.882	1509
$\widehat{R}_i\|_{L_i=(0,\frac{1}{2})}$	.658	.671	.671	.020	.612	.729	201
$\widehat{R}_i\|_{L_i=(\frac{1}{2},\frac{2}{3})}$	1.527	1.523	1.523	.017	1.477	1.550	130
$\widehat{R}_i\|_{L_i=(\frac{2}{3},\frac{4}{5})}$	2.833	2.836	2.842	.023	2.744	2.877	621
$\widehat{R}_i\|_{L_i=(\frac{4}{5},1)}$	6.734	6.649	6.650	.268	5.890	7.316	557
$\widehat{R}_i$	3.867^c	3.842^c	2.850	2.290	.612	7.316	1509

Source: own calculations
^{a} mean based on column (1) of Table 2. ^{b} statsitics based on ML estimates of model (2) but w/o life satisfaction and religiosity as covariates. ^{c} weighted average of the four preceding figures in same column, weights are the fractions of respondents in each response category.

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13/16 September, **Elias Braunfels**, “Further Unbundling Institutions”

14/16 September/April 2017, **Alexander W. Cappelen, Cornelius Cappelen, and Bertil Tungodden**», "False positives and false negatives in income distribution"

15/16 October, **Aline Bütikofer**, Eirin Mølland, and **Kjell G. Salvanes**, "Childhood Nutrition and Labor Market Outcomes: Evidence from a School Breakfast Program"

16/16 October**, Itziar Lazkano** and Linh Pham, "Do Fossil-Fuel Taxes Promote Innovation in Renewable Electricity Generation?

17/16 October, Kristiina Huttunen, Jarle Møen, and **Kjell G. Salvanes**, «Job Loss and Regional Mobility».

18/16 November**, Ingvild Almås, Alexander Cappelen, Bertil Tungodden**, «Cutthroat capitalism versus cuddly socialism: Are Americans more meritocratic and efficiency-seeking than Scandinavians?"

19/16 December, **Gernot Doppelhofer, Ole-Petter Moe Hansen**, Melwyn Weeks, "Determinants of long-term economic growth redux: A Measurement Error Model Averaging (MEMA) approach "

20/16 December, Leroy Andersland and **Øivind A. Nilsen**, "Households' responses to price changes of formal childcare"

21/16 December, Wilko Letterie and **Øivind A. Nilsen**, "Price Changes - Stickiness and Internal Coordination in Multiproduct Firms"

22/16 December, **Øivind A. Nilsen** and Magne Vange, "Intermittent Price Changes in Production Plants: Empirical Evidence using Monthly Data"

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01/17 January, **Agnar Sandmo**, "Should the marginal tax rate be negative? Ragnar Frisch on the socially optimal amount of work"

02/17 February, **Luca Picariello**, "Organizational Design with Portable Skills"

03/17 March, **Kurt R. Brekke**, Tor Helge Holmås, Karin Monstad og Odd Rune Straume, "Competition and physician behaviour: Does the competitive environment affect the propensity to issue sickness certificates?"

04/17 March, **Mathias Ekström**, "Seasonal Social Preferences".

05/17 April, Orazio Attanasio, Agnes Kovacs, and **Krisztina Molnar**: "Euler Equations, Subjective Expectations and Income Shocks"

06/17 April, **Alexander W. Cappelen**, Karl Ove Moene, Siv-Elisabeth Skjelbred, and **Bertil Tungodden**, "The merit primacy effect"

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09/17 June, **Astrid Kunze**, "The gender wage gap in developed countries"

10/17 June, **Kristina M. Bott, Alexander W. Cappelen, Erik Ø. Sørensen** and **Bertil Tungodden**, "You've got mail: A randomized field experiment on tax evasion"

11/17 August, Marco Pagano and **Luca Picariello**, "Talent Discovery, Layoff Risk and Unemployment Insurance"

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14/17 August, **Helge Sandvig Thorsen**, "The effect of school consolidation on student achievement".

15/17 September, Arild Sæther, "Samuel Pufendorf and Ludvig Holberg on Political Economy".

16/17 September**, Chang-Koo Chi,** Pauli Murto, and Juuso Välimäki**, "**War of attrition with affiliated values".

17/17 September, **Aline Bütikofer** and Giovanni Peri, "The Effects of Cognitive and Noncognitive Skills on Migration Decisions"

18/17 October, **Øivind Schøyen**, "What limits the powerful in imposing the morality of their authority?

19/17 October, **Charlotte Ringdal** and **Ingrid Hoem Sjursen**, "Household bargaining and spending on children: Experimental evidence from Tanzania"

20/17 December, **Fred Schroyen** and Karl Ove Aarbu, "Attitudes towards large income risk in welfare states: an international comparison"

Norges
Handelshøyskole

Norwegian School of Economics

NHH
Helleveien 30
NO-5045 Bergen
Norway

Tlf/Tel: +47 55 95 90 00
Faks/Fax: +47 55 95 91 00
nhh.postmottak@nhh.no
www.nhh.no