Risk-based Selection in Unemployment Insurance: Evidence and Implications

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Abstract

This paper studies whether adverse selection can rationalize a universal mandate for unemployment insurance (UI). Building on a unique feature of the unemployment policy in Sweden, where workers can opt for supplemental UI coverage above a minimum mandate, we provide the first direct evidence for adverse selection in UI and derive its implications for UI design. We find that the unemployment risk is more than twice as high for workers who buy supplemental coverage. Exploiting variation in risk and prices, we show how 25-30% of this correlation is driven by risk-based selection, with the remainder driven by moral hazard. Due to the moral hazard and despite the adverse selection we find that mandating the supplemental coverage to individuals with low willingness-to-pay would be sub-optimal. We show under which conditions a design leaving choice to workers would dominate a UI system with a single mandate. In this design, using a subsidy for supplemental coverage is optimal and complementary to the use of a minimum mandate.

Keywords: Adverse Selection, Unemployment Insurance, Mandate, Subsidy

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1 Introduction

While unemployment insurance (UI) systems vary in many dimensions across countries (e.g., the level and time profile of unemployment benefits), they share one striking similarity: employed workers are mandated to participate in UI and are not given any choice. They are forced to pay payroll taxes when employed and receive a set transfer when unemployed, which is not subject to choice. Why do (almost) all countries mandate UI? Why is no coverage choice available? Are these optimal features of UI design? Despite the large existing literature on UI, these fundamental questions have so far been unanswered.

A universal mandate is seen as the canonical solution to the inefficiencies arising under adverse selection [see Akerlof [1970], Chetty and Finkelstein [2013]]. Indeed, it is well-known that adverse selection hinders efficient market function as low risks leave the market and put upward pressure on equilibrium prices. While adverse selection is arguably the culprit in the context of UI, there are two issues with this argument. First, since UI is universally mandated, the role of adverse selection in UI cannot be directly tested. Second, even when adverse selection is present, the government may do better by using alternative interventions that allow for choice. Our paper tries to address both issues. We provide first-time evidence on the presence and severity of risk-based selection into unemployment insurance and we develop a general framework to evaluate the desirability of a universal mandate vs. choice-based interventions using this evidence.

We study this question in the Swedish context, where *all* workers are entitled to a minimum benefit level when becoming unemployed, but can opt to buy more comprehensive UI at a uniform premium set by the government. The combination of a mandate into basic coverage with a (subsidized) option for more generous coverage is used in other Scandanivian UI programs and common practice in other social insurance programs around the world.

We provide a theoretical framework - with both adverse selection and moral hazard - to evaluate the design of social insurance programs allowing for choice. The policy levers in our framework are the plan prices and coverage levels. When plan prices reflect the costs of individuals selecting those plans, the concern is that workers will 'under-insure'. In principle, both price and coverage levels can be adjusted - with a universal mandate as an extreme case - to induce people to buy more comprehensive coverage. The fiscal externality from steering workers from basic into comprehensive coverage equals the difference between the price and cost differentials for workers at the margin. This wedge will not only depend on how adversely selected plan choices are, but also on the moral hazard response of these workers to the extra coverage. The fiscal externality for the marginal workers needs to be compared to the welfare impact of the plan changes on the inframarginal workers. We derive sufficient-statistic formulae, combining insights from the Einav-Finkelstein [Einav et al. [2010b]] and Baily-Chetty [Baily [1978], Chetty [2006]] frameworks, that highlight the central trade-offs and can be implemented empirically. In particular, we demonstrate how, on the one hand, adverse selection in both the comprehensive and basic plan and, on the other hand, moral hazard among workers on either the comprehensive or basic plan, are essential inputs to the evaluation of the optimal price and coverage levels.

Our empirical analysis aims to provide estimates of these inputs, exploiting the combination of the exceptional setting and rich administrative data in Sweden. In particular, we observe the UI choice of the universe of Swedish workers and can link these choices to their unemployment histories registered by the Public Employment Service. We also merge this data with a rich collection of household and firm registers, providing extremely detailed information on the determinants of workers' unemployment risk and insurance choices. We present a set of empirical results, which provide direct and robust evidence that workers have private information about their unemployment risk, and act on this when making their unemployment insurance choice.

In a first step, we perform so-called positive correlation tests, assessing whether workers who choose to buy comprehensive UI are more likely to be unemployed [see Chiappori and Salanié [2000]]. Our estimates indicate that unemployment for workers buying the comprehensive coverage is about 2.3 times more likely than for workers who choose to stay on basic coverage. This large difference, however, reflects the combined impact of risk-based selection and moral hazard. To separate the two, we build a rich predictive model of individuals' *ex ante* unemployment risk, leveraging various features of the Swedish labor market that provide variation in unemployment risk beyond the direct control of individuals.[1] Our decomposition implies that the difference in *ex post* risk realizations driven by adverse selection is less than half of the wedge driven by moral hazard. Moreover, the moral hazard response to supplemental coverage is estimated to be large even for the workers who stick to basic coverage, unlike the "selection on moral hazard" findings in Einav et al. [2013] and Shepard [2016].

In a second step, we provide direct evidence of risk-based selection following an approach inspired by Einav et al. [2010b], which consists in using price variation to identify *marginal* buyers and compare their unemployment risk to *inframarginal* buyers of the same insurance plan. We contribute to the standard approach by offering a methodology, based on panel data, that allows for aggregate risk correlated with price variation. We exploit a large premium increase in 2007 and find evidence of significant risk-based selection: the unemployment risk for the workers at the margin, who stopped buying comprehensive coverage when the price increased, is $20-40\%$ higher than for the inframarginal workers who did not buy comprehensive coverage, neither before nor after the premium increase. Since their unemployment risk is measured under the same coverage, this difference cannot be driven by moral hazard. This difference in unemployment realizations between the marginal and inframarginal workers disappears when conditioning on predictable unemployment risk, which validates our earlier decomposition. In parallel to the price variation, we study how changes in benefits affect demand and risk-based selection. We exploit the cap on the unemployment benefit level of the comprehensive plan in a regression kink design. Among the unemployed workers, the share of workers on the comprehensive plan is increasing as the comprehensive benefit level is higher, but their average unemployment risk is going down, providing additional evidence

[1] We also use these risk shifters more directly to test for the presence of risk-based selection, similar in spirit to the unused observables test in Finkelstein and Poterba [2014].

of adverse selection.

In the final part of the paper we use our empirical estimates to evaluate a UI system with choice of coverage:

First, despite the severe adverse selection, our estimates indicate that adverse selection by itself cannot rationalize a universal mandate into comprehensive coverage in Sweden. The revealed value for workers who choose not to buy the comprehensive coverage is exceeded by the insurance costs. These costs are high due to the large estimated moral hazard response. As a result, mandating those workers to buy the comprehensive coverage would decrease welfare. This is of course an important conclusion in light of the universal mandates of comprehensive UI coverages in other countries and the absence of prior tests whether adverse selection can make such policy desirable.[2]

Second, the estimated adverse selection indicates an important role for subsidizing comprehensive coverage. Before the 2007 reform, the premium corresponded to only 31% of the difference between the respective average costs of providing the comprehensive and basic plan. The large subsidy encouraged around 86% of workers to buy the comprehensive plan. The 2007 price increase eliminated this subsidy, but the demand response has been relatively inelastic. Our analysis indicates that at the efficient price - at which the fiscal externality from encouraging workers to buy comprehensive coverage is zero - still 83% percent of workers would buy it. The high pre-reform subsidy, however, could still be rationalized by the redistributive gains away from workers on basic coverage towards workers on comprehensive coverage.

Third, the optimal coverage differentiation crucially depends on the difference in moral hazard costs. Our evidence suggests that for workers selecting the comprehensive coverage - who thus value the extra coverage more - the moral hazard cost from providing the extra coverage is actually lower than for workers selecting the basic coverage. This force suggests that maintaining a relatively large difference in UI benefits across plans can be optimal. However, we note that these conclusions are conditional on the level of prices: a decrease in the subsidy weakens the case for further coverage differentiation. Put simply, a generous minimum mandate and a large subsidy for comprehensive coverage are complementary policies.

Our work contributes to different strands of the literature. First, our work aims to contribute to a large and rapidly growing empirical literature analyzing the role of adverse selection in insurance markets [see for example Einav et al. [2010a]], by highlighting the advantages of using comprehensive, detailed and population-wide registry data and proposing new approaches to identify risk-based selection. Second, the lack of private markets and choices related to unemployment insurance, makes that the role of adverse selection in UI specifically has been untested so far. Most related to our paper is the work by Hendren [2017], who analyzes elicited beliefs about job loss and finds that workers' private information on their unemployment risk is sufficient to explain the absence of a private market for supplemental unemployment insurance in the US (in addition to the public UI policy in place). Our paper complements Hendren's evidence with direct evidence

[2]Examples of countries mandating UI with similar replacement rates as the voluntary, comprehensive plan in Sweden are Belgium, France, Luxembourg, Netherlands, Portugal, Spain and Switzerland. In other countries like the US and the UK, UI is also compulsory, but at lower replacement rates.

based on actual insurance choices and studies the optimality of the public unemployment policy itself. Finally, our work tries to bridge two strands of the social insurance literature, characterizing optimal coverage policies under moral hazard [Baily [1978], Chetty [2006], Schmieder et al. [2012], Kolsrud et al. [2018]] and characterizing optimal price policies under adverse selection [e.g., Hackmann et al. [2015], Tebaldi [2017], Finkelstein et al. [2019]]. Our conceptual framework provides implementable insights for policy design, related to recent work by Veiga and Weyl [2016] and Azevedo and Gottlieb [2017] who characterize equilibria with endogenous prices and coverages. In comparison, we explicitly allow for moral hazard and potential selection on moral hazard like in Einav et al. [2013] and Shepard [2016].

Our paper proceeds as follows. In Section 2 we describe the institutional background and the data we use. Section 3 introduces our theoretical framework. In Section 4 we provide positive correlation tests relating unemployment risk to UI coverage and decompose the positive correlation between adverse selection and moral hazard using predictable risk. In Section 5 we use price variation and benefit variation to provide evidence for risk-based selection and identify the statistics necessary to identify the welfare consequences of various policy interventions. Section 6 puts things together and determines the welfare impacts of various changes to the structure of the Swedish UI system. Section 7 concludes.

2 Context and Data

2.1 Institutional Background

Sweden is with Iceland, Denmark and Finland, one of the only four countries in the world to have a voluntary UI scheme, historically administered by trade union-linked funds (the so-called Ghent system). This is the system many countries had in place before switching to compulsory insurance overseen by the government [see Carroll [2005]]. The Swedish UI system consists of two parts:

The first part of the system is mandated and provides basic coverage funded by a payroll tax (that we denote p_0). The benefits that unemployed receive with this basic coverage (b_0) are non-contributory (i.e., do not depend on the unemployed earnings prior to displacement). The benefit level of the basic coverage is low. During our period of analysis (2002-2009) the benefit level remained at 320 SEK per day (≈35 USD) which corresponds to a replacement rate of a little less than 20% for the median wage earner.[3]

The second part of the Swedish UI system is voluntary. By paying an insurance premium $\mathbf{p} = p_1 - p_0$ to UI funds (on top of the payroll tax p_0), workers can opt for more comprehensive coverage. Upon displacement, workers who have continuously contributed premia for the comprehensive coverage during the past twelve months, get benefits b_1, that replace 80% of pre-unemployment earnings up to a cap, *in lieu* of the basic coverage b_0.[4] Apart from the benefit

[3]Benefits are paid per "working day", which means that there are 5 days of benefits paid per week. Benefits of 320 SEK a day therefore translate into 6960 SEK a month (≈765 USD).

[4]Enrolling in the supplemental coverage is done by filling out a form, which can be obtained online or in direct

level, there are no coverage differences between the basic and the comprehensive UI scheme. In particular, the potential duration of benefits b_0 and b_1 is the same, and was unlimited during our period of analysis. Moreover, to be eligible for either benefit upon unemployment, workers must fulfil a labor market attachment criterion, which is that they need to have worked 80 hours per month for six months during the prior year.[5]

The administration of the comprehensive UI coverage is done by 27 UI funds (so-called *Kassa's*) but the government, through the Swedish Unemployment Insurance Board (IAF), supervises and coordinates the entire UI system. In particular, both the premia and benefit levels of the basic and comprehensive coverage are fully determined by the government. To be clear, even though the funds are in charge of implementing the system, they are all operating under the rules set by the government, implying that Swedish UI is publicly provided.[6] Importantly, the government does not allow UI funds to charge different prices to different individuals. One exception are union members who get a small rebate of $\approx$ 10% on the UI premium for the comprehensive coverage, so in our empirical analysis, we always control for trade union membership to account for this.[7] During our period of study, the government also did not allow premia to differ across UI funds. Premia paid by workers cover only a (small) fraction of benefits paid by the UI funds to eligible unemployed, and the government subsidizes UI funds for the difference out of the general budget.

Until January 1st of 2007, the monthly premium $\mathbf{p}$ for the comprehensive coverage was homogeneous across UI funds, at around 100 SEK, and a 40% income tax credit was given for the premia paid. In January 2007, the newly elected right-wing government increased the premium substantially and removed the income tax credit on premia paid to UI funds. It also introduced an additional fee that partly tied the premium of each UI fund to the average unemployment rate of that fund, starting from July 2008. In our analysis, and partly due to data availability, we focus on the period before July 2008 where insurance premia are homogeneous across UI funds.

The combination of a mandate into basic coverage with a (subsidized) option for more generous coverage is not unique to the Swedish UI context, but commonly used in other social insurance programs including health insurance, old-age pensions, disability insurance, etc. In the US health

contact with the UI funds. The premium is paid monthly and enrolled members can select between receiving monthly invoices or paying via direct debit. In case the fee is not paid for three consecutive months, despite monthly reminders, the membership is terminated (the neglected payments must still be paid). A cheaper way to opt out of the plan is to fill out a form, analogous to the procedure of opting in. There are no waiting periods associated with opting in or out, and the processing time for such requests are typically limited to a few days.

[5] Note that the self-employed are given the same option to get comprehensive coverage. To actually receive UI benefits, they need to close their business [see Kolsrud [2018]]. In most countries, however, self-employed workers and the growing share of workers under alternative work arrangements are not covered by the UI system - either because they don't have access or are not mandated to participate [see OECD [2018]].

[6] Historically, with the "Ghent system" in place, labor and trade unions played an important role in providing unemployment insurance in Sweden. Today's 27 UI funds, which broadly correspond to 27 different industries/occupations, originated from unemployment insurance funds set up by unions. However, since the government overtook the responsibility of supervising the entire UI system in 1948, the links between UI funds and unions have loosened progressively.

[7] The 10% rebate on UI premia for union members is a remnant of the "Ghent system", but a large ($\approx$ 20%) and growing share of workers are members of an unemployment fund without being members of a union, and a growing share of union members ($\approx$ 10%) do not buy unemployment insurance. Note that individuals can still continue to contribute to UI funds while unemployed, for instance to build eligibility in case of a future unemployment spell, in which case they are also entitled to paying a reduced premium.

insurance market, for example, the recent Affordable Care Act involved the combined use of a minimum mandate and subsidies. Social security design often combines public pension benefits and tax-favored pension savings. To implement choice, a government may provide a menu of plan options by itself, or alternatively provide or mandate only basic coverage and count on private insurers to offer plans or top-ups.[8]

2.2 Data

We combine data from various administrative registers in Sweden. First, we use UI fund membership information for the universe of workers in Sweden aged 18 and above, from 2002 to 2009, and coming from two distinct sources. The first source is tax data for the period 2002 to 2006, during which workers paying UI premia received a 40% tax credit. The UI funds sent information annually to the Tax Authority about everyone who had contributed to the voluntary coverage plan within the year. Our data contain the total amounts of UI premia paid for each individual and year, as reported by the UI funds to the Tax Authority. From this source, we define a dummy variable D for buying the comprehensive coverage in year t as reporting any positive amount of premia paid in year t. For the analysis using the price variation of the 2007 reform in Section 5.1, we combine this data with a second source of information, coming from UI fund data that *Kassa's* sent to the IAF. This data contain a dummy variable indicating whether an individual aged 18 and above in Sweden is contributing premia for the comprehensive coverage as of December of each year from 2005 until 2009.

We add data on unemployment outcomes coming from the Swedish Public Employment Service, with records for the universe of unemployment spells from 1990 to 2015, and we merge it with the UI benefit registers from the IAF which provides information on all UI benefit payments (for both the basic and comprehensive coverage), information on daily wage for benefit computation, and Kassa membership information for all unemployed individuals. Based on this data, we define unemployment as a spell of non-employment, following an involuntary job loss, and during which an individual has zero earnings, receives unemployment benefits and reports searching for a full-time job. To define the start date of an unemployment spell, we use the registration date at the PES. The end of a spell is defined as finding any employment (part-time or full-time employment, entering a PES program with subsidized work or training, etc.) or leaving the PES (labor force exit, exit to another social insurance program such as disability insurance, etc.).[9] We define displacement as an involuntary job loss, due to a layoff or a quit following a 'valid reason'.[10] In the rest of the paper,

[8]See for example Cutler and Reber [1998] in the context of health insurance and Cabral and Cullen [2019] in the context of disability insurance.

[9]Note that UI benefits can be received *forever* in Sweden during the period 2002-2006 so the duration spent unemployed is identical to the duration spent receiving unemployment benefits.

[10]Valid reasons for quitting a job are defined as being sick or injured from working, being bullied at work, or not being paid out one's wage by one's employer. Quits are reviewed by the Public Employment Service at the moment an individual registers a new spell and if the quit is made because of a valid reason, the individual is eligible for UI and a notification is made in the PES data, allowing us to observe such quits under valid reasons. Involuntary quits are a small fraction of unemployment spells in our sample: 95.0% of unemployment spells observed in our data are due to layoffs. We exclude voluntary quits from our measure of unemployment and displacement.

we use the terms displacement and layoff as synonyms.

We complement the data with information on earnings, income, taxes and transfers and demographics from the LISA register, and with information on wealth from the wealth tax registers. We also exploit variation in unemployment risk across individuals due to Sweden's employment-protection law. In particular, we use the layoff-notification register (VARSEL) for years 2002 to 2012, which records the notifications by firms to the Public Employment Service as required by law when intending to displace 5 or more workers. The list needs to follow the last-in-first-out (LIFO) principle. For that, we use the matched employer-employee register (RAMS), from 1985 to 2015, to compute tenure and tenure ranking for for the universe of individuals employed in establishments of firms operating in Sweden.

2.3 Predictive Model of Unemployment Risk

We leverage the rich set of observables available in the Swedish registry data, and the various institutional features of the Swedish labor market to build a predictive model of unemployment risk. That is, the best predictor of future unemployment risk given all currently observed individual characteristics. This measure will allow us to go beyond studying how realized risk in year $t+1$ correlates with choice in t and also study how predictable risk in year t correlates with choices at time t. This will prove important in separating adverse selection from moral hazard.

Our main measure of unemployment risk π throughout the paper, and the one relevant to the UI system given insurance choices made in year t, is the number of days an individual is expected to spend unemployed in $t+1$.[11] To account for the fact that the distribution of days spent unemployed is defined only over non-negative integers, and exhibits a significant mass at zero, throughout the paper, we model π using a zero-inflated Poisson model. The expected number of days unemployed conditional on a vector of characteristics X therefore takes the following form:

$$E(\pi|X) = (1 - f(0|X_I)) \exp(X_C'\beta)$$

For the zero-inflated part of the process, we parametrize the probability $f(0)$ using a logit: $f(0|X_I) = \exp(X_I'\beta)/(1+\exp(X_I'\beta))$. We will allow the set of risk predictors X_I and X_C, entering respectively the inflated part and the count part, to differ.

The richness of the Swedish registry data allows us to observe many predictors of unemployment risk such as age, education, location, occupation, industry, earnings, etc. The Swedish institutional context also creates significant variation in unemployment risk that is arguably beyond the control of individuals. In Appendix A, we present evidence showing the importance of three risk shifters in particular, which will figure prominently in our vector X of risk predictors. The first risk shifter is the average (i.e. "leave-out mean") firm layoff rate. The second is layoff notifications: the risk

[11] If an individual has bought the comprehensive coverage throughout year t, then the days she spends unemployed in year $t+1$ will be covered by the comprehensive benefits. In that sense, the relevant risk to determine the cost of providing the comprehensive coverage to an individual buying that coverage in year t is the expected number of days she will spend unemployed in year $t+1$.

of unemployment increases significantly following layoff notifications. Finally, the enforcement of the Last-In-First-Out principle creates significant variation in unemployment risk within firm over time across individuals with different tenure levels.

In terms of model selection, we discipline the choice of the many potential regressors by using the adaptive Lasso procedure for a zero-inflated Poisson model proposed by Banerjee et al. [2018], that we detail in Appendix A.2. The regressors we allow to initially enter the model are individual log earnings, family type, nine age bins, gender, twelve dummies for education level, year fixed effects, region fixed effects, industry fixed effects, dummies for the past layoff history of the individual, dummies for the layoff notification history of the firm, the leave-out mean of firm layoff risk, union membership, tenure rank, interactions between tenure ranking and firm layoff risk and interactions between tenure ranking and layoff notification history of the firm. The Lasso procedure ends up mostly picking up the "institutional" risk shifters (i.e. layoff notification, tenure, etc.) in predicting displacement risk, while other demographics such as education or region also play an important role in the count part of the model. In Appendix A.2, we provide all further details on the estimation procedure.

To account for moral hazard, we allow the risk of individuals with similar characteristics X to differ if they are observed under the basic coverage or under the comprehensive coverage. To this purpose, we estimate separately two models of predicted risk. The first model is the predicted risk given X under the basic coverage $\hat{\pi}_0 = E(\pi_0|X)$. This model is estimated on individuals who are observed under the basic coverage in t. The second model is the predicted risk given X under the comprehensive coverage $\hat{\pi}_1 = E(\pi_1|X)$, which we estimate on individuals who are observed under the comprehensive coverage in t.

To assess the quality of the model fit, Figure 1 shows bin scatters of the relationship between predicted risk under basic (resp. comprehensive) coverage and actual realized risk for individuals under basic (resp. comprehensive) coverage. In both panels, the relationship is close to the 45-degree line indicating that the model does a good job at predicting the average realization of unemployment risk. However, the model slightly under-predicts very long unemployment spells for workers under comprehensive coverage (see Panel B). Comparing both panels, we also see that individuals under basic coverage have lower realized unemployment risk, and thus lower predicted risk than individuals in the comprehensive coverage. We provide additional elements of diagnostics on the quality of our model fit and summary statistics on the distribution of predicted risk in Appendix A.2. In general, we find significantly less dispersion in our predicted measure of risk than in realized risk. This confirms that there still remains a substantial dimension of idiosyncratic unemployment risk beyond what can be predicted even using a very rich set of observables.

2.4 Summary Statistics

In Table 1, we characterize the empirical setting by providing summary statistics for our main sample of interest over the period 2002 to 2006. The sample consists of individuals aged between 18 and 60 and who have been working for at least 6 months. The average probability to be displaced in year $t+1$ conditional on working in year t is 3.0% over the period 2002 to 2006. The average probability to be unemployed in year $t+1$ (unconditional on employment status in year t) is higher, at 3.6%. The average (unconditional) number of days unemployed in $t+1$ is 5.28. Workers in our sample are predicted to spend on average 3.57 days in $t+1$ if under the basic coverage and 5.83 days if under the comprehensive coverage. Note also that the fraction of individuals who are members of a UI fund (i.e., buying the comprehensive UI coverage) is large during the 2002-2006 period, at 86%. The Table shows that there is also limited switching over time across coverages over the period 2002 to 2006. In Appendix Table B.3 we provide further summary statistics breaking down the sample between individuals observed under the basic coverage and individuals observed under the comprehensive coverage. The Table shows that individuals under the basic coverage are younger, are more likely to be men and to be single, and hold significantly larger wealth and liquid assets than individuals under comprehensive coverage.

3 Conceptual Framework

This section presents a conceptual framework that accounts for adverse selection and moral hazard and underpins our empirical and welfare analysis. We first set up a model of UI choice in Sweden, where a minimum benefit level is mandated, but workers can opt for comprehensive coverage. We then use the framework to characterize the key trade-offs when setting prices and coverages as a function of estimable moments. A universal mandate - with no choice offered - can be considered as an extreme case of setting prices and/or coverages such that all individuals are on the same plan.[12]

3.1 Setup

Workers are offered the choice between two plans that differ in the coverage they provide against unemployment risk: a basic plan (b_0, p_0) and a comprehensive plan (b_1, p_1). They can opt for a higher UI benefit level $b_1 \geq b_0$, but this comes at a higher price $p_1 \geq p_0$. The coverages and prices are the levers of the government's unemployment policy. These policy levers affect workers' selection of plans and their unemployment risk. The setup encompasses a universal mandate, when $(b_0, p_0) = (b_1, p_1)$ and no choice is allowed for.

The key micro-foundations for our analysis are workers' plan valuations and the government's costs and how both change with the plans' prices and coverage levels. Worker i chooses the plan providing the highest utility $u_i(b_j, p_j)$. She will thus opt for the comprehensive plan when

$$u_i(b_1, p_1) \geq u_i(b_0, p_0). \tag{1}$$

We will use short-hand notation u_1, u_0 and $\mathbf{u} = u_1 - u_0$ respectively. A worker's unemployment risk depends on her type and the actions she undertakes given her coverage. For tractability, we assume that workers' preferences are quasi-linear in prices so that an individual's risk, conditional

[12] See Appendix F for further discussion and proofs.

on plan choice j, does not depend on prices and neither does the ranking of individuals' valuations $\mathbf{u}_i$. Individual i's unemployment risk under coverage b_j is denoted by $\pi_i(b_j)$ (or π_j in short). The average unemployment risk for workers who opt for coverage j if they are under plan j' equals

$$E_j\left(\pi_{j'}\right) = E\left(\pi_i\left(b_{j'}\right)|u_i\left(b_j, p_j\right) \geq u_i\left(b_{-j}, p_{-j}\right)\right). \tag{2}$$

The worker's unemployment risk determines the cost to the government of providing coverage, denoted by $c_i(b_j) = \pi_i(b_j) b_j$.

3.2 Social Insurance Design

Our aim is to characterize how to set prices and coverages when both adverse selection and moral hazard are present. Both forces have been the subject of large, but surprisingly parallel literatures in social insurance. As is well known, adverse selection makes it inefficient to price insurance plans at average cost, while moral hazard makes it inefficient to provide complete coverage.

We assume a concave and differentiable social welfare function,

$$\mathcal{W} \equiv \int_{\mathbf{u}_i \geq 0} \omega\left(u_i\left(b_1, p_1\right)\right) di + \int_{\mathbf{u}_i < 0} \omega\left(u_i\left(b_0, p_0\right)\right) di + \lambda\left\{F_1\left[p_1 - E_1\left(\pi_1\right) b_1\right] + F_0\left[p_0 - E_0\left(\pi_0\right) b_0\right]\right\},$$

where the function $\omega(\cdot)$ maps individuals' utility into social welfare, λ equals the marginal cost of public funds, which pre-multiplies the fiscal cost of the unemployment policy, and F_j denotes the share of individuals buying plan j for given coverages and prices.

Our main focus is on the fiscal externalities of workers' choices and how they change with prices and coverages. We ignore the presence of other frictions or inefficiencies, but revisit later the potential role of choice frictions due to behavioral biases [see Spinnewijn [2017]] and the ex-ante value of insurance [see Hendren [Forthcoming]], which can both drive a wedge between the welfare-relevant utility and the decision utility at the time a decision is made.

PCT Decomposition We first provide two complementary ways to quantify the respective roles of adverse selection and moral hazard, which relate directly to the fiscal externalities from changing the prices and coverage levels as derived below. Both adverse selection and moral hazard increase the correlation between unemployment risk and coverage, $E_1(\pi_1) - E_0(\pi_0)$. A first decomposition of this PCT statistic is into the difference in risks for the two groups under comprehensive coverage and the difference in risks under the two plans for the group selecting basic coverage,

$$E_1\left(\pi_1\right) - E_0\left(\pi_0\right) = \underbrace{E_1\left(\pi_1\right) - E_0\left(\pi_1\right)}_{AS_1} + \underbrace{E_0\left(\pi_1 - \pi_0\right)}_{MH_0}. \tag{3}$$

The former term captures adverse selection into the comprehensive plan (AS_1), while the latter term captures moral hazard for the group selecting basic coverage (MH_0). The alternative decomposition is into moral hazard for the group selecting comprehensive coverage (MH_1) and adverse selection into the basic plan (AS_0),

$$E_1(\pi_1) - E_0(\pi_0) = \underbrace{E_1(\pi_1 - \pi_0)}_{MH_1} + \underbrace{E_1(\pi_0) - E_0(\pi_0)}_{AS_0}. \tag{4}$$

Differences in moral hazard among the individuals on comprehensive and basic coverage relate mechanically to differences in adverse selection in the comprehensive and basic plan. Appendix Figure F.1 provides a graphical illustration of these relations, linking this to the textbook analysis of selection and treatment effects.[13]

Price Policy We now characterize the impact of a price change dp_j on social welfare. We consider a small deviation, so that we can invoke the envelope theorem: the impact on individuals' welfare depends on the direct effect of the policy change, but not on the behavioral response and re-sorting of individuals at the margin. The direct welfare effect of a price change dp_j on the buyers of plan j depends on the marginal social value of income for these workers, for which we use short-hand notation $E_j\left(\frac{\partial \omega_j}{\partial p_j}\right)$. This welfare effect should be compared to the fiscal impact of the price change, which depends on both the direct revenue change dp_j for the share of (inframarginal) workers buying plan j and on the fiscal externality of the share of (marginal) workers switching in or out of comprehensive coverage.

The fiscal externality due to the selection response depends on the difference between the price differential $\mathbf{p} = p_1 - p_0$ and cost of providing comprehensive instead of basic coverage to the marginal buyers. This corresponds to the well-known result in Einav et al. [2010b]. Denoting the unemployment risk of the marginal buyers by $E_{M(\mathbf{p})}(\pi_j) = E(\pi_j|\mathbf{u} = 0)$, we obtain

$$FE_{\mathbf{p}}^{AS} \equiv [p_1 - p_0] - \left[E_{M(\mathbf{p})}(\pi_1) b_1 - E_{M(\mathbf{p})}(\pi_0) b_0\right], \tag{5}$$

$$= \left[E_1(\pi_1) - E_{M(\mathbf{p})}(\pi_1)\right] b_1 + \left[E_{M(\mathbf{p})}(\pi_0) - E_0(\pi_0)\right] b_0 - S, \tag{6}$$

where $S = [E_1(\pi_1) b_1 - E_0(\pi_0) b_0] - [p_1 - p_0]$ denotes the subsidy for supplemental coverage capturing how much the price differential differs from the average cost differential.

Equation (6) demonstrates how in our binary choice setting the fiscal externality accounts for risk-based selection in both the comprehensive and basic plan. When both plans are priced at average cost and thus $S = 0$, adverse selection typically causes the fiscal externality to be positive. To be more precise, if the marginal buyer is less risky than the average buyer of the comprehensive coverage and more risky than the average buyer of the basic coverage, the government will gain twice from inducing this marginal buyer to switch from basic to comprehensive coverage.[14] Approximating

[13] The differences in AS and MH generally relate to topics of heterogeneous treatment effects and selection into treatment [e.g. Kowalski [2016]; Kline and Walters [2019]]. So-called selection on moral hazard [Einav et al. [2013]] can be interpreted as $MH_1 - MH_0 > 0$. The opposite can happen as well, but requires the difference in risks under basic coverage to be larger than the difference in risks under comprehensive coverage, $AS_0 - AS_1 > 0$. Indeed, $MH_1 - MH_0 = AS_1 - AS_0$ immediately follows from decompositions (3) and (4).

[14] The expression for the fiscal externality in equation (6) also shows that selection and not moral hazard itself drives the inefficiency of average-cost pricing. The impact of moral hazard on the cost differential would be priced

$E_1(\pi_1) - E_{M(\mathbf{p})}(\pi_1) \approx F_1 \times [E_1(\pi_1) - E_0(\pi_1)]$ and $E_{M(\mathbf{p})}(\pi_0) - E_0(\pi_0) \approx F_0 \times [E_1(\pi_0) - E_0(\pi_0)]$, the fiscal externality can be linked to the adverse selection terms in the PCT decompositions, (3) and (4),

$$FE_{\mathbf{p}}^{AS} \approx F_1 AS_1 b_1 + F_0 AS_0 b_0 - S, \tag{7}$$

confirming that the fiscal externality depends on risk-based selection in both the comprehensive and basic plan.[15]

Comparing the welfare effects from changing prices of the respective plans, we can state:

Proposition 1. *For given coverage levels, the prices p_0 and p_1 are optimal only if*

$$\frac{E_1\left(\frac{\partial\omega_1}{\partial p_1}\right)}{E_0\left(\frac{\partial\omega_0}{\partial p_0}\right)} = \frac{1 + FE_{\mathbf{p}}^{AS}\frac{\partial \ln F_1}{\partial p_1}}{1 - FE_{\mathbf{p}}^{AS}\frac{\partial \ln F_0}{\partial p_0}}.$$

The left-hand side of Proposition 1 captures the redistributive gain from transferring a marginal dollar from individuals on the basic plan to those buying the comprehensive plan. The right-hand side equals the fiscal return of these transfers due to the change in plan selection. For example, a reduction in the premium for comprehensive coverage induces more workers to buy it and the return to this selection effect is positive as long as the fiscal externality is positive ($FE_{\mathbf{p}}^{AS} > 0$). In the absence of redistributive motives, the optimal subsidy is such that the fiscal externality equals zero ($FE_{\mathbf{p}}^{AS} = 0$). The price differential then exactly reflects the cost of providing the supplementary coverage to individuals at the margin. Increasing the subsidy further would cause the fiscal externality to be negative ($FE_{\mathbf{p}}^{AS} < 0$), but could be justified by valuing redistribution from workers on basic coverage towards workers on comprehensive coverage.

Coverage Policy We now turn to the impact of a change in coverage db_j. An increase in coverage provides more insurance to the group of workers selecting this plan, but also reduces their incentives to avoid unemployment. This standard trade-off between insurance and incentives is captured by the well-known Baily-Chetty formula [Baily [1978], Chetty [2006]] applied to the workers selecting a given plan.

Like in the Baily-Chetty formula, the fiscal externality of providing extra coverage depends on the moral hazard response by these workers, captured by the increase in their unemployment risk as coverage increases,

$$FE_{b_j}^{MH} = E_j\left(\frac{\partial \pi_j}{\partial b_j}\right)\frac{b_j}{E_j(\pi_j)}. \tag{8}$$

However, a key difference with the standard Baily-Chetty formula comes from the extra fiscal gain

efficiently under average-cost pricing if the moral hazard impact were constant across workers.

[15] The approximation would be exact under rank-linearity $E(\pi_j|\mathbf{u}) = \alpha_j + \beta_j G(\mathbf{u})$ for $G(\cdot)$ the cdf. This for example holds when demand and cost curves are linear as in Einav et al. [2010b].

or cost due to the selection response to a coverage change. Like for a price change, we need to account for the share of switchers into or out of comprehensive coverage. The corresponding fiscal externality equals

$$FE_{b_j}^{AS} \equiv [p_1 - p_0] - \left[E_{M(b_j)}(\pi_1)\, b_1 - E_{M(b_j)}(\pi_0)\, b_0\right], \tag{9}$$

where $E_{M(b_j)}(\pi_k) = E\left(\pi_k \frac{\partial u_j}{\partial b_j} | \mathbf{u} = 0\right) / E\left(\frac{\partial u_j}{\partial b_j} | \mathbf{u} = 0\right)$ is a weighted average of the unemployment risk among the marginal buyers under plan k. In comparison with the fiscal externality of a price change $FE_{\mathbf{p}}^{AS}$, higher weight is given to the risk of the marginal buyers who value the extra coverage more as they are more likely to switch.[16] Of course, the coverage levels are no stand-alone instruments and can be used in combination with subsidies. For example, setting a more generous basic coverage level b_0 will worsen the adverse selection into comprehensive coverage, but the worsened adverse selection can be addressed with a more generous subsidy. In particular, a higher subsidy will reduce the corresponding fiscal externality in equation (9).[17]

Comparing the welfare effects from changing coverages of the respective plans, we can highlight the value of differentiating the coverages among which workers can choose:

Proposition 2. *For given prices, the coverage levels b_0 and b_1 are optimal only if*

$$\frac{E_1\left(\frac{\partial \omega_1}{\partial b_1}\right) / E_1(\pi_1)}{E_0\left(\frac{\partial \omega_0}{\partial b_0}\right) / E_0(\pi_0)} = \frac{1 + FE_{b_1}^{MH} - FE_{b_1}^{AS} \frac{\partial \ln F_1}{\partial b_1} / E_1(\pi_1)}{1 + FE_{b_0}^{MH} + FE_{b_0}^{AS} \frac{\partial \ln F_0}{\partial b_0} / E_0(\pi_0)}.$$

The left-hand side of Proposition 2 equals the ratio of the insurance gain from increasing the coverage of the comprehensive vs. the basic plan. The value from extra coverage for workers on either plan, $E_j\left(\frac{\partial \omega_j}{\partial b_j}\right)$, depends on their marginal value from extra UI when unemployed.[18] The right-hand side equals the relative fiscal cost of increasing the coverages, depending on both the moral hazard responses and the selection responses discussed above. The value of differentiating the coverage levels b_1 vs. b_0 thus comes from the fact that individuals who value extra coverage can opt for it. By revealed preference, we expect individuals who opt for extra coverage to value it more, but the returns to differentiation are decreasing as workers are risk averse. The cost of differentiating the coverage levels depends on how high the moral hazard cost is among workers selecting comprehensive vs. basic coverage, but also on the fiscal return to encouraging more individuals to opt for comprehensive coverage.

[16] The differential selection depending on plan characteristics has been studied for example in Veiga and Weyl [2016], but also relates to the difference in LATE's depending on the instruments used [e.g., Kline and Walters [2019]; Mogstad et al. [2019]]. We show this formally in Appendix F.1. Note that if workers differ only along a one-dimensional index, the marginal buyers responding to a change in coverage or in price would be the same, as would the corresponding fiscal externalities.

[17] The worsening selection in response to a generous minimum mandate has been studied before in Azevedo and Gottlieb [2017] (see also Finkelstein [2004] and Chetty and Saez [2010]), but also provides a different perspective on the absence of private UI (see Hendren [2017]), which is conditional on the mandated public UI that is already in place.

[18] With expected utility and utilitarian social welfare, the scaled value term $E_j\left(\frac{\partial \omega_j}{\partial b_j}\right) / E_j(\pi_j)$ simplifies to the average marginal utility of consumption when unemployed, just like in the standard expressions of the Baily-Chetty formula, but now for the workers on plan j.

Just like for the adverse selection externality, we can link the moral hazard externality back to our earlier PCT decompositions,

$$FE^{MH}_{b_j} \approx MH_j \times \frac{b_j}{[b_1 - b_0]E_j(\pi_j)}.$$

Here we approximate the relevant marginal moral hazard response using the unemployment risk response to a switch between comprehensive and basic coverage. This approximation indicates that so-called selection on moral hazard (i.e., $MH_1 > MH_0$), where the moral hazard response is larger for workers on comprehensive coverage, would weaken the argument for more differentiation. Risk-based selection, however, either in the comprehensive plan (AS_1) or in the basic plan (AS_0), would increase the fiscal return from inducing workers to switch to comprehensive coverage and tends to strengthen the argument for more differentiation in coverage levels.

Universal Mandate Our analysis sheds light on the value of offering choice more generally. The most common policy in the context of UI is to impose a universal mandate, not allowing for any choice. The key question is whether introducing choice is desirable when starting from a universal mandate. Or alternatively, when starting from a differentiated schedule, whether less differentiation in coverages is desirable. Proposition 2 identifies the moments that allow answering this question and helps deriving a simple non-parametric test to evaluate whether a universal mandate into one of the coverage levels would be desirable. For example, we can evaluate the welfare gains from a universal mandate into b_1 by considering the corresponding coverage increase for the workers under basic coverage. In line with the proposition, this is simply the sum of the insurance gains net of moral hazard costs from the incremental coverage increases going from the basic to comprehensive coverage, highlighting again that a potential impediment to mandating workers into comprehensive coverage is moral hazard among those who value the comprehensive coverage the least. Alternatively, we can consider a (sufficiently) large increase in p_0 (or decrease in p_1) such that all workers under the basic plan switch to the comprehensive plan. In terms of efficiency consequences, the conclusions are exactly the same. In line with Proposition 1, the efficiency gain from such a universal mandate is simply the sum of the fiscal externalities $AS_{\mathbf{p}}$ corresponding to the required marginal price changes. Following Einav et al. [2010b], this corresponds to the valuation of the supplemental coverage to the workers under the basic coverage relative to the cost of providing it. Using a standard revealed preference argument, we can bound the valuation of the supplemental coverage for these workers from above by the price differential $p_1 - p_0$ (which they are not willing to pay). Hence, the earlier PCT decomposition in (3) allows for a simple, non-parametric test for the desirability of a universal mandate. Mandating all workers into comprehensive coverage is inefficient if

$$\begin{aligned} p_1 - p_0 &\leq b_1 E_0(\pi_1) - b_0 E_0(\pi_0), && (10) \\ &= (b_1 - b_0) E_0(\pi_0) + b_1 MH_0. && (11) \end{aligned}$$

This test again underlines the importance of the moral hazard among the buyers of basic coverage, but it does set the redistributive consequences aside and also assumes the absence of other frictions that may justify a universal mandate.

Empirical Implementation. In the next two sections we turn to the empirical identification of the various AS and MH terms determining the desirability of a mandate, and the welfare consequences of changes to the price and benefit of UI policies in the Swedish context. We proceed in three steps. First, we start with positive correlation tests and propose a decomposition of the test statistic using predictable risk, separating the AS and MH terms. This allows, under some testable assumption, for the identification of the desirability of a mandate. We then use variation in price. This allows to identify AS terms directly, and enables the validation of our decomposition between AS and MH terms. With this evidence, the welfare consequences of changes to the price structure can be evaluated. We finally focus on benefit variation, which enables the identification of demand responses to coverage levels and of AS for individuals at the margin of benefit variation. With this evidence, the welfare consequences of changes to the benefit structure can be evaluated.

4 Positive Correlation Tests

A natural first step to investigate adverse selection is to produce correlation tests. We therefore start by showing the presence of a strong positive correlation between an individual's choice of UI coverage and her unemployment risk. But correlation tests cannot disentangle the respective role of adverse selection and moral hazard. Using our predicted risk model, we then show the presence of a positive correlation between UI choices and predictable risk. These correlations confirm the presence of significant adverse selection in both the basic ($AS_0 > 0$) and comprehensive coverage ($AS_1 > 0$). We then use predictable risk to propose a decomposition of the positive correlation between selection and moral hazard, the validity of which depends on a testable assumption. This decomposition also allows for the identification of MH_0, the moral hazard cost created by individuals who would be moved to the comprehensive coverage by a mandate.

4.1 Positive Correlation Tests in Realized Risk

The correlation test consists in comparing the expected risk of individuals conditional on their insurance coverage choice. In particular, we test for $E_1(\pi_1|Z) > E_0(\pi_0|Z)$, where the vector Z controls for characteristics that affect the unemployment insurance contracts available to an individual.[19] Over our baseline period of interest (2002-2006), UI contracts only differ according to three dimensions.

[19] Controlling for these characteristics guarantees that we compare individuals who are facing the same options so that the correlation is driven by demand rather than by supply (different individuals being offered different contracts by the *Kassa*). As explained in Section 2.1 above, characteristics affecting the premia and benefits under each coverage are strictly regulated by the government.

The first dimension is employment history. Coverage depends on whether individuals meet a work eligibility requirement or not, for which they need to have worked for at least 6 calendar months within the past 12 months prior to displacement. We therefore include in vector Z an indicator for having worked at least 6 months in year t.[20]

The second dimension of contract differentiation is earnings. As explained in section 2.1, due to the presence of a benefit cap, the additional daily benefits $\mathbf{b} \equiv b_1 - b_0$ that individuals get when buying the supplemental coverage is a kinked function of daily earnings w. Formally, $\mathbf{b} = F(w) = (.8 * w - 320) \cdot \mathbb{1}[400 \leq w < 850] + 360 \cdot \mathbb{1}[850 \leq w]$. We therefore include the supplemental benefit function $F(w)$ as a control function in Z to make sure that we compare individuals facing the same benefit level per unit of premium paid.

The last dimension of contract differentiation is that union members pay a slightly lower premium than non-union members for the supplemental coverage. We therefore include in Z an indicator variable for union membership. We also include year fixed effects in Z to account for small adjustments to the premium in January every year over the period 2002-2006.

To test in practice for $E_1(\pi|Z) > E_0(\pi|Z)$, we use our baseline measure of risk π, which is the total duration (in days) spent unemployed in year $t+1$. And we correlate this measure of risk with insurance choices made in year t. In practice, we estimate the following zero-inflated Poisson process specification:

$$E(\pi|Z) = (1 - f(0|Z, D)) \exp(Z'\beta + \alpha \cdot \mathbb{1}[D = 1]), \tag{12}$$

where $D \equiv \delta(\mathbf{u} \geq 0)$ is an indicator for buying the comprehensive coverage.[21] We estimate specification (12) on the pooled sample of all individual $i \times$ year t observations between 2002 and 2006.

The first bar of Figure 2 reports the semi-elasticity of days unemployed in $t+1$ with respect to insurance choice in t, estimated from model (12):

$$\text{Semi}_{PCT} = \frac{E(\pi|Z, D = 1) - E(\pi|Z, D = 0)}{E(\pi|Z, D = 0)}. \tag{13}$$

Results indicate a strong and significant positive correlation between realized risk and UI coverage choice: Individuals who buy the comprehensive coverage in t spend 135% more days in unemployment in $t+1$ than individuals who stick to the basic coverage in t. Our results are robust to the use of alternative measures of unemployment risk and to functional form specifications, as shown in Appendix B.

[20] Note that eligibility requires individuals to have worked at least 80 hours per month for 6 calendar months within the past 12 months. While we do not have precise data on monthly hours, to be conservative, we also include a dummy for having earnings above 80 hours $\times$ 6 months $\times$ the negotiated janitor wage. In the absence of an official, legally binding minimum wage in Sweden, the janitor wage is often considered the effective minimum wage in the labor market.

[21] As explained above, the bolded notation $\mathbf{u}_i$ refers to the difference in expected indirect utility between plan 1 and plan 0 for individual i.

4.2 Selection on Predictable Risk

The predicted risk model presented in section 2.3 allows us to test how much UI coverage choices correlate with predictable risk at time t, rather than realized risk at $t+1$. This approach is akin to the "unused observables" test of Finkelstein and Poterba [2014], and provides evidence that the PCTs are not just driven by moral hazard but also by risk-based selection.

In the spirit of that test, we first gauge how the selection into coverage depends on specific institutional risk shifters, which enter the predicted risk model and are arguably beyond the control of individuals. In Appendix C, we show that UI coverage choice is strongly correlated with average firm layoff risk in the cross-section. Moreover, using a firm switcher design, we find that the probability to buy comprehensive UI increases significantly when moving to a firm with a higher turnover risk. Finally, within a firm, workers are more likely to start buying the comprehensive coverage when the layoff risk increases as proxied by the firm's sending of a layoff notification to the PES, and this effect is strongest among individuals with lower relative tenure within occupation$\times$establishment cells, as predicted by the application of LIFO rules. The identifying variation and affected workers differ for the three strategies, but the large and significant responses of UI coverage choice indicate significant risk-based selection into UI.

By using the predicted risk model, we can leverage the variation in predictable risk in a comprehensive way and study the corresponding adverse selection. We use our baseline sample over the period 2002-2006, and follow a specification similar to (12). We start by using as an outcome the risk measure $\hat{\pi}_0$, which corresponds to the unemployment risk (in days) that an individual is predicted to face in $t+1$ given her characteristics in t, were she to be under the basic coverage in t. Results are presented in Figure 2. The second bar of the graph reports the semi-elasticity of $\hat{\pi}_0$ with respect to insurance choice defined in (13). It reveals that the group of individuals observed choosing the comprehensive coverage in t are predicted to have $\approx 30\%$ more days unemployed in $t+1$ than individuals who do not buy, *if both groups were hypothetically observed under the same basic coverage.*

We then turn to using $\hat{\pi}_1$ as an outcome, which corresponds to the unemployment risk predicted under the comprehensive coverage. The third bar of Figure 2 reports the semi-elasticity of $\hat{\pi}_1$, with respect to insurance choice in t. We find again a strong and significant positive correlation between insurance choice and predicted risk. We note that the semi-elasticity of both measures of predicted unemployment duration $\hat{\pi}_0$ and $\hat{\pi}_1$ are significantly smaller than the semi-elasticity for realized unemployment duration in $t+1$ (first bar of Figure 2). As explained below, this difference can be explained by the presence of significant moral hazard.[22]

4.3 Decomposition of PCT between Selection and MH

Under the assumption that there is no residual unobserved risk correlated with UI choice, all relevant adverse selection in the basic and in the comprehensive coverage is identified by the difference in

[22] Appendix C provides further non-parametric evidence on the relationship between predicted risk and insurance choice.

predictable risk between individuals observed in the comprehensive and in the basic coverage. Formally, under the assumption that $E_1[\pi_1|\hat{\pi}_1] = E_0[\pi_1|\hat{\pi}_1]$ (i.e., conditional on the predicted risk, the average risk under comprehensive coverage is the same for both groups), selection in the comprehensive coverage is:

$$AS_1 \equiv E_1(\pi_1) - E_0(\pi_1) \;=\; E_1(\hat{\pi}_1) - E_0(\hat{\pi}_1). \tag{14}$$

Equivalently, under the assumption that $E_1[\pi_0|\hat{\pi}_0] = E_0[\pi_0|\hat{\pi}_0]$ then selection in the basic coverage is:

$$AS_0 \equiv E_1(\pi_0) - E_0(\pi_0) \;=\; E_1(\hat{\pi}_0) - E_0(\hat{\pi}_0). \tag{15}$$

Importantly, the assumption underpinning (14) and (15) can be tested. In section 5.2, we use price variation to identify willingness-to-pay, and we validate that there is no residual variation in risk correlated with willingness-to-pay, when conditioning on our predicted risk measure.

Based on (14) and (15), we can get estimates of adverse selection into comprehensive (AS_1) and basic coverage (AS_0) from our predicted risk model, and then use these estimates to provide decompositions of the PCT, between selection and moral hazard. Following formula (3), we can decompose the PCT between AS_1 and moral hazard for individuals selecting into the basic coverage MH_0. Alternatively, we can decompose the PCT between AS_0 and MH_1, i.e., moral hazard for individuals selecting into the comprehensive coverage, using formula (4).

We implement these decompositions in Figure 3 using our main sample over the period 2002-2006, when on average 86% of individuals buy comprehensive coverage.

The decomposition shows large differences in predicted risk under comprehensive vs. basic coverage. Despite the presence of significant adverse selection, most of the positive correlation between risk and insurance choices is driven by moral hazard. The share equals 62 and 75 percent respectively for the two alternative decompositions. The exercise also indicates the presence of some small selection on moral hazard (i.e., $MH_1 > MH_0$), but that conclusion is reversed when expressing the unemployment risk responses proportionally to the risk under basic coverage for the respective groups (which is how the moral hazard terms enter the welfare characterizations in Proposition 2).[23] As stated before, the decompositions rely on the assumption that our predicted risk model absorbs all variation in risk correlated with willingness-to-pay. If for instance, there is adverse selection in the residual unemployment risk conditional on predictable risk, then, our exercise will provide a lower bound on adverse selection, and an upper bound on moral hazard. We now turn to using price and coverage variation to identify selection and we also use this to validate our decomposition.

[23]The corresponding moral hazard elasticities are respectively .77 for individuals under comprehensive, and .94 for individuals under basic coverage. The finding of substantive moral hazard is in line with the large literature estimating the elasticity of unemployment durations/exit rates with respect to unemployment benefits: Schmieder and von Wachter [2016] summarize estimates from 18 studies from 5 different countries, and find a median of estimate of 0.53. Kolsrud et al. [2018] find an even larger elasticity of 1.53(.13) in the Swedish context.

5 Identifying Risk-based Selection using Plan Variation

In this section we exploit variation in both price and coverage levels to provide direct evidence on risk-based selection, building on Einav et al. [2010b]. The selection effects determine the fiscal externality of price and coverage interventions, following the welfare framework of Section 3. We also use the price variation to test and validate the assumption underlying our earlier decomposition between moral hazard and adverse selection.

5.1 The 2007 Price Reform

We first exploit a sudden and unanticipated increase in the premia paid to get the supplemental coverage in 2007. The reform followed the surprise ousting of the Social Democrats from government after the September 2006 general election. With this reform, monthly premia, which had been remarkably stable over the previous years, suddenly increased from 100 SEK to around 320 SEK on January 1st, 2007, as shown in Figure 4. The Figure also shows that the take-up of comprehensive coverage responded significantly to this sharp surge in prices. After staying almost constant around 86%, the fraction of the eligible population buying the comprehensive coverage abruptly dropped to 78% right after the reform. Interestingly, Figure 4 displays little sign of pre-trends or anticipation in the take-up rate of the comprehensive coverage, adding credibility to the assumption that this sudden increase in premia, following the surprise change in political majority, was arguably exogenous to individuals' willingness-to-pay for the comprehensive coverage. The unemployment rate was also smoothly decreasing throughout the period, so that the increase in $\mathbf{p}$ cannot be explained by an endogenous pricing response to an increase in the underlying costs of the comprehensive coverage.[24]

5.2 Risk-Based Selection Response to Prices

The 2007 reform created significant variation in price and in the fraction buying the comprehensive coverage. Following Einav et al. [2010b], this variation could be exploited to identify adverse selection by simply comparing average costs of providing comprehensive coverage across the different price levels, i.e. before vs after the reform. Yet, in our context, variation in average costs may also reflect realizations of some aggregate unemployment risk, which may vary over time, and will therefore correlate with the price variation. More generally, if there is some aggregate component to risk, and if there is correlation between aggregate risk variation and price variation, direct comparisons of average costs across price observations as in Einav et al. [2010b] will not identify adverse selection. This can be an issue in insurance contexts, where most of the variation in price

[24] If anything, the 2007 premia reform was combined with a minor legislated *decrease* in the benefits received in the comprehensive coverage. On January 1st 2007, the cap on the benefits b_1 was slightly decreased for benefits received in the first 20 weeks of an unemployment spell. Given this reform had only a negligible effect on average benefits received, we neglect it in the welfare implementation.

available comes from variation over time, or across places and groups of individuals.[25]

We propose a simple method to address this issue and identify adverse selection. We use the fact that with panel data, exogenous price variation allows for the identification of marginals, who switch coverage in response to the price change. We can then rank individuals in three groups, ordered in terms of their valuation of the supplemental coverage ($\mathbf{u}$). First, we define group 1 as the group of individuals insured in the comprehensive coverage both in 2006 and 2007: they were buying the supplemental coverage in 2006 under the low premia and continue to buy the supplemental coverage under the high premia, and therefore have the highest valuation of the comprehensive coverage ($\mathbf{u} > 0$). We then define the group of marginals $M(\mathbf{p})$, who were buying the comprehensive coverage in 2006 but switch out in 2007 when premia $\mathbf{p}$ increase: these individuals have a lower willingness-to-pay for supplemental insurance than individuals from group 1 and are close to indifferent between the two coverages at current prices ($\mathbf{u} \approx 0$). Finally, individuals who were neither buying the comprehensive coverage in 2006 nor in 2007, and are therefore always under the basic coverage are defined as group 0: they have the lowest willingness-to-pay for the supplemental coverage ($\mathbf{u} < 0$).[26]

Using this ranking, we can now perform direct non-parametric tests for risk-based selection, by correlating willingness-to-pay with measures of unemployment risk π. Because the marginals and the individuals from group 0 are now observed under the same basic UI coverage, the comparison of the average realized risk of these two groups under basic coverage ($E_{M(\mathbf{p})}(\pi_0) - E_0(\pi_0)$) is immune to moral hazard and provides a direct estimate of risk-based selection.[27] Because individuals are compared under the same aggregate conditions, this test is also immune to aggregate risk realizations correlated with the price variation.[28]

Figure 5 presents the results of such non-parametric tests and provides direct evidence of the presence of risk-based selection into UI. Panel A starts by reporting the average number of days spent unemployed in 2008 for each group. We condition again on the vector Z of controls for contract differentiation, similar to what we did in the positive correlation tests. We define the variable D as taking value 1 for group 1, M for marginals, and 0 for group 0. We then estimate

[25]For example, Hackmann et al. [2015] use price variation over time in the context of health insurance using a difference-in-differences design.

[26]Note that our partition of the population ignores a negligible fourth group of individuals, who were not buying the comprehensive plan in 2006, but switched in the comprehensive plan in 2007. The size of this group is seven times smaller than the group of individuals switching out of the comprehensive plan in 2007. The reason we exclude this group of workers from the analysis is that their ranking in terms of willingness-to-pay is ambiguous, as we discuss in Appendix D.

[27]It is worth re-emphasizing the timing of the Swedish UI policy: one needs to contribute for at least 12 months in order to become eligible to the comprehensive benefits b_1. Marginals and individuals from group 0 in 2007 did not contribute any premium to the comprehensive plan in 2007. In 2008, if they become unemployed they will therefore get the basic benefits b_0 irrespective of their insurance choice in 2008. In other words, because of their insurance choice in 2007, marginals and group 0 individuals face the exact same coverage in 2008. The difference in their unemployment risk in 2008 cannot be driven by moral hazard due to different coverage choices in 2008.

[28]In a similar spirit, Shepard [2016]compares the costs of individuals staying and switching out of a plan in response to a change in plan characteristics in the year *before* the change, when both groups were under the same plan.

specification:

$$E_j(\pi|Z) = (1 - f(0|Z, D)) \exp(Z'\beta + \sum_j \alpha_j \mathbb{1}[D = j]). \tag{16}$$

For each panel of Figure 5, we report $E(\pi|Z = \bar{Z}_0, D = j)$, the average realized risk outcome π of each group $j = 1, M, 0$, evaluated at the average value of Z for group 0.

Panel A shows that the average realized unemployment risk in 2008 of the marginals is significantly higher (22%) than that of individuals who are always under the basic coverage, while both groups are eligible to the same coverage in 2008. This is direct evidence of risk-based selection. That is, a positive correlation between risk and willingness-to-pay for the supplemental coverage. We can also see in panel A that there is a large and significant difference in the average realized risk of the marginals and individuals from group 1. Because individuals from group 1 and the marginals are now observed under different coverages, this difference identifies $E_1(\pi_1) - E_{M(\mathbf{p})}(\pi_0)$ and is a combination of selection and moral hazard.

In the last two panels of Figure 5, we report the relationship between willingness-to-pay and predictable risk, using our predictive model of days spent unemployed. Panel B plots the average predicted risk under basic coverage $E_j(\hat{\pi}_0|Z)$ for the three groups $j \in \{1, M, 0\}$, where we use the same method as in panel A to control for the vector Z of characteristics affecting contract differentiation. Similarly, panel C plots the average predicted risk under comprehensive coverage $E_j(\hat{\pi}_1|Z)$ for the same three groups. Note that in both panels, we use the risk $\hat{\pi}_0$ and $\hat{\pi}_1$ predicted by our model using individuals' observable characteristics as of 2006.[29] Comparing the predictable risk of marginals individuals ($j = M$) to individuals with the lowest valuation of the comprehensive coverage ($j = 0$), we find in both panels B and C the presence of significant adverse selection.

In Appendix D, we provide further results and probe into the robustness of our results. In particular, we investigate the robustness to using alternative measures of risk and address potential concerns, such as inertia, to the validity of our ranking of individuals by willingness-to-pay. We also provide additional results showing that our ranking of willingness-to-pay for the comprehensive coverage correlates strongly with determinants of the insurance value and proxies for risk preferences.

Validation of PCT decomposition The evidence leveraging the price variation shows the presence of significant adverse selection, but also suggests the presence of sizeable moral hazard. For each group of workers - even those with the lowest willingness-to-pay for insurance - the predicted unemployment risk is much larger in the comprehensive coverage than in the basic coverage, as shown in the last two panels of Figure 5. Like for the decomposition of the PCT in section 4.3, the separation between adverse selection and moral hazard requires that, once we condition on our measure of predicted risk, there is no residual unobserved heterogeneity in risk correlated with insurance choices. The price variation offers the possibility to test this assumption directly.

[29] We fix observable characteristics as of 2006, prior to the price change, as individuals might have changed these characteristics endogenously in 2007 based on their new insurance coverage choice, which would reintroduce potential moral hazard. Fixing observable characteristics as of 2007 instead gives nevertheless very similar results.

Because we observe the risk under the basic coverage π_0 of groups $j = M$ and $j = 0$, we can test for $E_{M(\mathbf{p})}(\pi_0|\hat{\pi}_0) = E_0(\pi_0|\hat{\pi}_0)$.

Figure 6 displays the results. The left bar in the graph starts by reporting the difference in realized risk in 2008 for the marginals ($j = M$) and for the individuals always in the basic coverage ($j = 0$) when simply controlling for the vector of observables Z affecting the contract space. To control for these observables, we use specification (16) above, and report on the graph:

$$\text{Semi}_{M(\mathbf{p})}^{Baseline} = \frac{E(\pi|Z, D = M) - E(\pi|Z, D = 0)}{E(\pi|Z, D = 0)},$$

This is the estimated semi-elasticity of the average realized risk under basic coverage for the marginals M relative to the individuals always under basic coverage in 2006 and 2007.

To determine whether any correlation between risk and willingness-to-pay survives when controlling for predicted risk, we now compute the residual semi-elasticity:

$$\text{Semi}_{M(\mathbf{p})}^{Residual} = \frac{E_{M(\mathbf{p})}(\pi_0|Z, \hat{\pi}_0) - E_0(\pi_0|Z, \hat{\pi}_0)}{E_0(\pi_0|Z, \hat{\pi}_0)},$$

where we condition on predictable risk by including in specification (16) twenty dummies for the ventiles of predicted unemployment risk under basic coverage in both the inflated and count part of the model.

Results, displayed in Figure 6, show that the semi-elasticity $\text{Semi}_{M(\mathbf{p})}^{Residual}$ drops to a tightly estimated zero when adding predicted risk as a control so that we cannot reject that $E_{M(\mathbf{p})}(\pi_0|\hat{\pi}_0) = E_0(\pi_0|\hat{\pi}_0)$. This evidence indicates that there is no significant residual correlation left between realized risk and willingness-to-pay when we fully control for predicted unemployment risk using the rich set of predictors from our predicted risk model. In other words, conditional on our predicted risk score, if there remains any unobservable idiosyncratic component to risk, it is uncorrelated with willingness-to-pay.

This result is particularly interesting. It suggests that little private information is left once we condition on this very detailed set of proxies for unemployment risk. In the second column of Figure 6, we investigate how much private information would be left if instead of controlling for our predicted risk score, we controlled non-parametrically for a rich set of observable demographic characteristics. We do so by including in specification (16) a set of dummies for age, gender, marital status, education (four categories), industry (1-digit code) and wealth level (quartiles). Interestingly, the semi-elasticity increases compared to our baseline when including these controls. This suggests that these characteristics provide advantageous selection on average, such that if contracts were differentiated along these observable dimensions, adverse selection into comprehensive coverage would actually be more severe.[30]

[30] An important point to note is that when layoff decisions are mandated by rules at the govt level like the LIFO principle in Sweden, those rules can affect the extent to which unemployment risk depends on observable vs. unobservable charateristics. A corollary of this point is that the above result that observables capture all dimensions of risk correlated with insurance choices in Sweden may not generalize to other labor markets like the US where firms

5.3 Regression Kink Design in UI Benefits

We now turn to variation in coverage levels and investigate the presence of adverse selection along the benefit margin. To obtain benefit variation, we leverage the kinked schedule of UI benefits b_1 in the comprehensive plan as a function of pre-unemployment earnings. Workers in the comprehensive plan receive daily unemployment benefits equal to 80% of their daily wage prior to unemployment, up to a cap. Over the period 2002 to 2007, this cap in daily UI benefits was fixed at 680SEK, meaning that the relationship between b_1 and daily wage w exhibited a kink at $w = 850SEK$.[31]

We identify the effect of variation in b_1 on demand and risk using a RK design, taking advantage of the kinky schedule of b_1 as a function of the daily wage. Our identifying variation is displayed in Figure 7 panel A, which plots a bin-scatter of the relationship between the daily wage and the observed average replacement rate for unemployed individuals in year t who have bought the comprehensive coverage in $t-1 \in [2002; 2006]$. The replacement rate is computed as the average benefit received during unemployment from the IAF data divided by the daily wage. The graph shows first that the replacement rate for b_1 is close to exactly 80% on the left hand side.[32] The graph also displays a clear kink at $w = 850SEK$, with the replacement rate declining sharply, as benefits are capped. We use this kinked relationship and treat it as a fuzzy RKD around the 850SEK threshold.

To implement this RK design, we use a measure of the daily wage coming from the IAF data. Unfortunately, this measure is only available for individuals who become unemployed, and for whom the IAF has to systematically collect this information to determine daily UI benefits. There does not exist a corresponding measure of the daily wage for the universe of workers in the registry data. For this part of the analysis, we therefore restrict our sample to individuals who become unemployed in year $t \in [2003; 2007]$ and were employed in year $t-1$. For these individuals, we relate their daily wage in year $t-1$ to their insurance choice in $t-1$ as well as their predictable risk in $t-1$. A consequence of this sample restriction is that individuals in this sample have a higher average risk than in our original sample, as riskier individuals in $t-1$ are more likely to be observed unemployed in t (see Appendix Table E.2). We note however that this does not affect the internal validity of our quasi-experimental analysis within this sample.

The key identifying assumption of the RK design is the existence of a smooth relationship at the threshold $w = 850SEK$ between the assignment variable and any pre-determined characteristics affecting the demand for insurance. We assess the credibility of this assumption by conducting two types of analysis in Appendix E. Considering the probability density function of the assignment variable, we do not detect lack of smoothness around the kink that would indicate the presence of selection. However, we find limited selection along observable characteristics around the kink, but show robustness of our results when controlling for the corresponding vector of characteristics X.

have more flexibility in determining who they let go.

[31] A daily wage of 850SEK corresponds to about 468USD a week using the average exchange rate over the period 2002 to 2007 of 1SEK $\approx$ 0.11USD.

[32] Note that the reason why the replacement rate is slightly below 80% is that some workers have their UI benefits reduced due to sanctions.

Like for the price variation, we first show that the demand for comprehensive coverage responds positively to the level of UI benefits b_1 offered in the comprehensive plan. Our RK estimand of the demand response is given by:

$$\delta = \frac{\lim_{w^-} dE[D|w]/dw - \lim_{w^+} dE[D|w]/dw}{\lim_{w^-} dE[b_1|w]/dw - \lim_{w^+} dE[b_1|w]/dw}, \tag{17}$$

where D is, as in section 4 above, an indicator variable for buying the comprehensive coverage. We estimate the numerator of the estimand based on the following specification:

$$D = \beta_0 \cdot (w-k) + \beta_1 \cdot (w-k) \cdot \mathbf{1}[w > k] + Z'\gamma_Z + X'\gamma_X, \tag{18}$$

where the coefficient β_1 corresponds to the numerator in (17). The vector Z controls, as before, for the characteristics that affect the premium paid for the comprehensive coverage over the period 2002 to 2006. It includes a dummy for union membership, a dummy for eligibility based on past employment history, and year fixed effects.

Figure 7 panel B provides a graphical representation of our results, using a bandwidth of 350SEK. It plots the average fraction of individuals who bought the comprehensive coverage in a given year, by bins of daily wage in that year. The relationship exhibits a sharp and significant downward kink at the $w = 850SEK$ wage threshold, which mirrors the sharp downward kink in the relationship between daily wage and benefits in panel A. This indicates that the demand for the comprehensive coverage responds negatively to a decline in the level of benefits b_1 in that coverage. We report on the graph the corresponding estimate of $\beta_1 * 100 = -.016(.001)$ along with its robust standard error. Based on the Delta-method, we obtain from this estimate a 95 % confidence interval for the elasticity of comprehensive coverage with respect to benefits $\frac{\partial 1-G(0)}{\partial b_1} \cdot \frac{b_1}{1-G(0)} \in [.13; .19]$.

5.4 Risk-Based Selection Response to Coverages

We now test for the presence of adverse selection along the benefit margin. Like with adverse selection along the price margin, we expect the marginals, who opt out of the comprehensive coverage as benefits b_1 decrease, to have lower risk π_1 in the comprehensive coverage than the individuals who stay in the comprehensive coverage: $E_1(\pi_1) > E_{M(\mathbf{b})}(\pi_1)$. Hence, a direct test for adverse selection in the comprehensive coverage à la Einav et al. [2010b] is therefore that the average risk of individuals in the comprehensive coverage goes up in response to the decrease in demand for comprehensive coverage at the kink. Formally, we test for a significant positive change in slope at $w = k$ in the relationship between the daily wage and the average risk in comprehensive coverage of the individuals observed in that coverage.

To implement this test, we conduct a RK analysis using our predicted risk measure $\hat{\pi}_1$ under comprehensive coverage, which we first residualize on the same vectors Z and X as in (18).[33] Panel

[33] As in sections 4 and 5 above, we use a zero-inflated Poisson structure to model the role of covariates Z and X on risk $\hat{\pi}_i$, $i \in \{0, 1\}$:

$$E(\hat{\pi}_i|Z, X) = (1 - f(0|Z, X)) \exp(Z'\gamma_Z + X'\gamma_X).$$

D of Figure 7 plots the relationship between the daily wage and the average residual risk measure $\tilde{\hat{\pi}}_1 = \hat{\pi}_1 - E(\hat{\pi}_1|Z, X)$ among individuals in the comprehensive coverage.[34] The panel suggests the presence of adverse selection: the average predicted risk in the comprehensive coverage is somehow flat as a function of daily wage on the left-hand side of the kink, but does increase significantly on the right hand side, kinking upwards around $w = 850$. To formalize the test, we collapse $\tilde{\hat{\pi}}_1$ at the daily wage bin level, and run the following RK specification:

$$\tilde{\hat{\pi}}_1 = \beta_0 \cdot (w - k) + \beta_1 \cdot (w - k) \cdot \mathbf{1}[w > k] \tag{19}$$

using the number of observations per bin as analytical weights. We report the estimate of $100 * \beta_1 = .279$ (.109) on the panel, which confirms the existence of a significant positive change in slope. We obtain from these estimates a 95 % confidence interval for the elasticity of the average predicted risk in the comprehensive coverage w.r.t to b_1: $\frac{\partial E[\hat{\pi}_1]}{\partial b_1} \cdot \frac{b_1}{E[\hat{\pi}_1]} \in [.07; .48]$.

These results indicate that individuals do adversely select in the comprehensive coverage in response to benefit variation. The average cost of providing the comprehensive coverage is therefore decreasing as we expand the share of individuals in that coverage through increasing benefits b_1, which is important for evaluating the welfare impact of changing coverages, following Proposition 2.

We can replicate this analysis for the average risk under the basic coverage. Risk-based selection implies that marginals switching into the basic coverage at the kink are riskier than individuals who were already in the basic coverage: $E_{M(\mathbf{b})}(\pi_0) > E_0(\pi_0)$. Panel C of Figure 7 plots the relationship between the daily wage and the average residual risk measure $\tilde{\hat{\pi}}_0 = \hat{\pi}_0 - E(\hat{\pi}_0|Z, X)$ among individuals in the basic coverage. Despite the lack of precision, due to the much more limited number of individuals observed in the basic coverage, the graph does indicate again the presence of an upward kink with an estimated change in slope: $100 * \beta_1 = .204$ (.144).

In Appendix E, we further investigate the robustness of our results. We show that our estimates are stable across bandwidth choices. We also document that the inclusion of the controls X has little effect on the overall results. Finally, we perform permutation tests for inference, using placebo kinks à la Ganong and Jäger [2018], and confirm the presence of significant adverse selection.

6 Policy Implications

This section uses our empirical estimates to evaluate the policy instruments introduced in Section 3. In particular, the Swedish setting allows us to test for the first time whether adverse selection can rationalize a generous UI mandate, which is commonplace in other countries. We also leverage our estimates to evaluate the fiscal externalities from changing the price and coverage levels.

[34] For readability, we rescale the residual by the average predicted risk at the average values for covariates Z and X of individuals at the kink ($w = k$). So formally, the panel plots: $\hat{\pi}_1 - E(\hat{\pi}_1|Z, X) + E(\hat{\pi}_1|\bar{Z}_{w=k}, \bar{X}_{w=k})$.

Universal mandate The high replacement rate in the comprehensive plan in Sweden is comparable to the replacement rates of UI mandates in many other countries, especially in Europe. As shown in Section 3, we can test for the desirability of mandating all workers into the comprehensive coverage by comparing the price of the extra coverage to its cost for workers choosing basic coverage. To that purpose, we obtain cost estimates by converting the predicted days spent unemployed (in Figure see Figure 5) using $E_0(c_j) = E_0(\hat{\pi}_j) \cdot b_j$ for both plans $j = 1, 0$.[35] The difference between the two gives an estimate for the cost of providing the extra coverage of $1,712 SEK$ to workers choosing the basic plan. This estimate reflects the large predicted moral hazard response when these workers opting for basic coverage would be put under comprehensive coverage. In comparison, before the 2007 reform in Sweden, workers had to pay a net premium of $720 SEK$ per year, accounting for the 40 percent tax credit. By revealed preference, this provides an upper bound on the valuation of workers opting for basic coverage, which is nevertheless substantially below the corresponding costs. As these workers value the coverage less than it costs, their choice not to buy comprehensive coverage is efficient. Imposing a universal mandate that forces them to buy the comprehensive coverage would be inefficient. This observation, reflecting the strong moral hazard response for workers with low valuation, raises the question whether a universal mandate of generous UI coverage observed in many other countries is socially desirable.

Setting Prices While a universal mandate may not be desirable, prices can still be used to overcome adverse selection by encouraging more workers into comprehensive coverage. Before the price increase in 2007, comprehensive coverage was heavily subsidized with the premium paying for only 31% of the difference in average costs. To calculate the fiscal externality form inducing workers to buy comprehensive coverage instead of basic coverage, following Proposition 1, we construct cost curves in the spirit of Einav et al. [2010b]. As discussed in Section 5, the price increase itself identifies individuals at the margin between the two plans and allows us to study how the cost of providing coverage changes with willingness-to-pay.

To construct cost curves, we convert the predicted risks (see Figure 5) into cost estimates, using $E_j(c_{j'}) = E_0(\hat{\pi}_{j'}) \cdot b_{j'}$ for the three groups of individuals $j = 1, M, 0$ and for both plans $j' = 1, 0$. We then position the three groups on the x-axis according to their willingness-to-pay, as shown in Panel A of Figure 8. Individuals who choose the basic coverage (0) both in 2006 and 2007 have the lowest valuation and are on the right hand side, while individuals who always buy the comprehensive coverage (1) are on the left-hand side of the graph. The marginals correspond to the group in between, and their share is given in Figure 4, where we see the fraction of individuals on comprehensive coverage dropping from 86 to 78% with the 2007 price increase. We then use a linear extrapolation to obtain marginal cost curves for providing the comprehensive and basic coverage respectively.[36] Reflecting the earlier results on predicted risks, the cost of the marginals

[35] To account for the taxes paid on the unemployment benefits received, we scale down the costs by .20, which corresponds to the empirical average tax rate paid by the unemployed (see Kolsrud et al. [2018]).

[36] We locate the estimated costs for the three groups at the midpoint of the ranges of the corresponding valuation quantiles and then use a piece-wise linear interpolation to construct the cost curves. Appendix Figure F.2 plots

is in between the cost for those with higher and lower valuation, both under comprehensive and under basic coverage.

We can now evaluate the fiscal externality from inducing workers to switch from basic to comprehensive coverage, i.e., the difference between the price and cost of providing supplemental coverage. This is shown in Panel B of Figure 8. The black triangles correspond to the difference in costs from providing the comprehensive and basic coverage, shown in Panel A. Providing the extra coverage is on average more costly for the workers who value it more, but the extrapolation suggests that for the group of switchers the cost stays between $1,929SEK$ and $2,051SEK$.[37] The grey triangles show the pre- and post-reform prices and the corresponding take-up of comprehensive coverage. By revealed preference, the valuation of the extra coverage for the group of switchers is bounded between the pre-reform price of $720SEK$ and the post-reform price of $4,116SEK$. The difference between revealed value and cost is negative for the marginal workers at the pre-reform price, indicating that is was efficient for them to buy comprehensive coverage. The opposite is true at the post-reform price. The fiscal externality is thus expected to be zero for some price in between. Assuming a linear demand curve, as plotted in Panel B, the premium at which value and cost coincide equals $2,022SEK$. 83 percent of workers would buy insurance at this efficient price.

While our estimates thus suggest that the large pre-reform subsidy induced too many individuals to buy comprehensive coverage, setting the subsidy as high can be rationalized by valuing the redistribution towards workers buying the comprehensive coverage. Evaluating the fiscal externality per marginal worker $FE^{AS}_{\mathbf{p}}$ at the pre-reform price, the right-hand side of Proposition 1 equals

$$\frac{1 + FE^{AS}_{\mathbf{p}} \frac{\partial \ln F_1}{\partial p_1}}{1 - FE^{AS}_{\mathbf{p}} \frac{\partial \ln F_0}{\partial p_0}} = 1.29,$$

using estimated price elasticities based on the demand response to the 2007 price increase (e.g., $\frac{\partial \ln F_0}{\partial \ln p_0} = .12$, see Figure 4). Hence, the large pre-reform subsidy would be optimal if the return from redistributing from workers under basic coverage to workers under comprehensive coverage equals 29%.[38]

Setting Coverages In Sweden the minimum UI benefit equals $320SEK$ ($\approx 35USD$) a day, which is about half of the average benefit level under the comprehensive plan. Proposition 2 characterized the welfare impact from further differentiating these coverage levels, showing that

the corresponding risk curves under comprehensive and basic coverage and shows how comparable the realized vs. predicted risks are for the group of marginal buyers under basic coverage.

[37]Note the slope of this marginal cost curve - capturing the cost of providing supplemental coverage - does no longer provide a test for adverse selection as in Einav et al. [2010b]. E.g., this curve can be upward sloping, while both the cost curve under comprehensive coverage and under basic coverage are downward sloping.

[38]Hendren [Forthcoming] proposes an approximation of the redistributive gain, given by $E_1\left(\frac{\partial \omega_1}{\partial p_1}\right)/E_0\left(\frac{\partial \omega_0}{\partial p_0}\right) \approx 1 + \frac{u''(c)}{u'(c)} \times [\mathbf{p} - E_0(\tilde{\mathbf{p}})]$, where the right-hand side depends on the curvature of utility in consumption and the average willingness-to-pay for comprehensive coverage for those under basic coverage, $E_0(\tilde{\mathbf{p}})$. Using the linear demand curve, the implied redistributive gain is less than 5% for $CARA = 5 \times 10^{-4}$ (expressing values in USD) or for $CRRA = 3$ (for an average annual consumption level of $150,000SEK$).

the corresponding fiscal externality depends on two major forces.

The first force is the relative magnitude of the moral hazard costs, MH_{b_1} and MH_{b_0}, depending on the unemployment risk response to a marginal change in the benefit level of workers who are in the comprehensive and basic coverage respectively. We measure these two terms using the arch-elasticity approximation:

$$FE_{b_j}^{MH} \approx MH_j \times \frac{b_j}{[b_1 - b_0]E_j(\pi_j)}.$$

where MH_j is our estimate of the moral hazard response to a change in benefit from b_0 to b_1 for individuals under coverage j. We find that: $FE_{b_0}^{MH} = .94 > FE_{b_1}^{MH} = .77$. In other words, the moral hazard externality of increasing benefits at the margin is lower for individuals under the comprehensive coverage than for individuals under the basic coverage. This force therefore pushes in favor of differentiating the benefit coverages.

The second force shaping the fiscal externality in Proposition 2 comes from the selection responses. In an adversely selected market, we expect these selection responses to strengthen the argument for benefit differentiation by decreasing the average risk of workers in both plans. Our estimates in section 5.4 indeed suggest that an increase in the comprehensive coverage attracts more workers in the comprehensive plan and reduces simultaneously the average costs of the comprehensive and of the basic plans. (A decrease in the benefit level of the basic coverage is expected to do the same.) However, the fiscal externality depends not only on the risk-based selection responses, but also on the level of the subsidy for the comprehensive coverage (cf. formula (6)). When the subsidy is high, the fiscal externality can be small or even negative despite significant adverse selection, as the cost of providing the extra subsidy outweighs the gains from the decrease in the costs of providing the coverages.

Evaluated again at the pre-reform price, we find that the fiscal externality of increasing the coverage in the comprehensive plan relative to the basic plan equals:

$$\frac{1 + FE_{b_1}^{MH} - FE_{b_1}^{AS}\frac{\partial \ln F_1}{\partial b_1}/E_1(\pi_1)}{1 + FE_{b_0}^{MH} + FE_{b_0}^{AS}\frac{\partial \ln F_0}{\partial b_0}/E_0(\pi_0)} \approx \frac{1 + .77 + .06}{1 + .94 - .51} = 1.27.$$

For this computation, we assume (i) that the demand response to a change in basic coverage is the same as the response to a change in comprehensive coverage and (ii) that the switchers are the same when considering variations in p, b_0 or b_1.[39,40]

The fiscal externalities in Proposition 2 need to be traded off against the relative consumption

[39] The first assumption can be formally stated as: $\frac{\partial F_1}{\partial b_1}/E_M(\pi_1) \approx \frac{\partial F_0}{\partial b_0}/E_M(\pi_0)$, evaluated for workers at the margin. This approximation would be exact under risk-neutrality.

[40] The second assumption always holds when the heterogeneity across agents can be captured by a one-dimensional index. The reason for relying on the second assumption is that we unfortunately lack power to credibly infer the predicted risk for the marginal buyers with respect to the coverage levels from the RKD estimates of the change in demand and average costs, which are both imprecisely estimated. Assuming $FE_{b_1}^{AS} \approx FE_{b_0}^{AS} \approx FE_{\mathbf{p}}^{AS}$, we can use our fiscal externality estimates from the price variation exercise of section 5.

smoothing gains from extra coverages: $\frac{E_1\left(\frac{\partial \omega_1}{\partial b_1}\right)/E_1(\pi_1)}{E_0\left(\frac{\partial \omega_0}{\partial b_0}\right)/E_0(\pi_0)}$. Further differentiation would therefore be socially valuable if the marginal utility of an increase in benefits for individuals under the comprehensive coverage is worth at least 27% more than the marginal utility of an increase in benefits for individuals under the basic coverage. A back-of-the-envelope calculation based on recent estimates from Landais and Spinnewijn [Forthcoming] in the same Swedish UI context suggests that the value of extra coverage would be around twice as high for workers under comprehensive coverage compared to workers under basic coverage, even when evaluated at b_1 and b_0 respectively.[41] This would imply that further differentiation in coverage levels enhances welfare and that providing the right choice to workers also dominates a UI system with a single mandate.

Finally, we note that the welfare implications of coverage differentiation are conditional on prices. At the current prices, the fiscal externality of inducing further individuals to switch to the comprehensive coverage is negative because the subsidy is too generous. By the same token, the selection response to a more generous minimum mandate would generate a positive fiscal externality. Instead, when the subsidy were to be lowered, the resulting changes in the adverse selection externalities would strengthen the case for further differentiation of the coverage levels.

Frictions and Timing of Choice Our welfare analysis has assumed that workers' willingness-to-pay for extra coverage reveals their valuation. We briefly discuss how choice frictions and the timing of choice can affect the above policy implications.

First, choice frictions drive a wedge between workers' willingness-to-pay and their welfare-relevant valuation (Spinnewijn [2017]). For example, the workers choosing not take the comprehensive coverage may well underestimate its value. The price ($720SEK$) is only just above the mechanical cost of providing the extra coverage ($686SEK$), not accounting for the cost increase due to the unemployment response (i.e., $E_0(\pi_0)[b_1 - b_0]$). The pricing is thus close to - and for the workers with lower risk better than - actuarially fair. One potential explanation why these workers still choose not to buy is that they underestimate their unemployment risk or simply underestimate the coverage they would get from comprehensive coverage.[42] Another explanation is that these workers over-estimate the price by not fully accounting for the 40% tax credit, but focus on the before-tax premium of $1200SEK$ instead. If the revealed preference approach makes us under-estimate the value of comprehensive coverage, the optimal subsidy should be higher and a universal mandate may become desirable. Our findings, however, still imply that in Sweden the severe adverse selection by itself is not sufficient to rationalize making the generous comprehensive

[41] In Appendix F we provide further detail on the evidence in Landais and Spinnewijn [Forthcoming] and the corresponding back-of-the-envelope calculation, which crucially depends on two forces: heterogeneity in the value of insurance, conditional on risk, and diminishing marginal utility of consumption. People who buy the comprehensive coverage b_1 may do so because they value an additional kroner of consumption when unemployed more, conditional on risk. However, diminishing marginal utility of consumption makes the value of an additional kroner lower when evaluated at b_1 than at b_0.

[42] See Landais and Spinnewijn [Forthcoming] for more evidence and discussion on the role of risk perceptions in this context, which is, however, inconclusive on the sign of the wedge between willingess-to-pay and welfare-relevant valuation.

coverage compulsory. Moreover, given the large moral hazard response, the choice frictions have to be substantial to change that conclusion.

Second, the presence of frictions may affect the estimated responses to changes in prices and coverage.[43] The 2007 reform increased the premium dramatically and, in fact, set it above the difference in average cost of providing comprehensive and basic coverage respectively. While the implied subsidy was thus eliminated, more than 3 out of 4 workers continued to buy comprehensive insurance. Interestingly, this suggests that the inelastic demand could hold this adversely selected market together when privatizing this market.[44] The inelastic response may, however, be due to inertia [e.g., Handel [2013], Polyakova [2016]]. To gauge this further, we study the insurance choice of workers who switch firms and arguably face a more active choice environment. Appendix Figure D.5 shows a comparable drop in demand for the comprehensive plan in the year after the premium increase among firm switchers. The decrease in take-up extends to the next year among workers who switch employers, while it flattens out for those who do not. We also investigated differences in adverse selection for workers in a more vs. less active choice environment. Appendix Figure D.6 shows that the job switchers who switched out of comprehensive coverage in 2007 are again more risky than the job switchers who were already in basic coverage in 2006. The semi-elasticities are of comparable magnitude for those who switched employers and those who did not. While this does not exclude the potential role of other frictions, it suggests that for the large 2007 reform inertia has played a small role.

Third, related to choice frictions, the timing of choice may affect welfare calculations as well (see Hendren [Forthcoming]). This will be particularly important in the presence of aggregate risk, as this affects whether workers' willingness-to-pay and their costs are sufficient to evaluate welfare. The extent to which aggregate risk affects the identification of the welfare-relevant curves depends on the timing of choice, that is, on the amount of aggregate risk that is realized at the time individuals make their choice. From the cost side, if aggregate unemployment risk is not realized by the time that individuals choose to buy insurance, using ex post realized costs rather than expected cost at the time of insurance choice would bias the welfare conclusions. In our implementation, to proxy for the expected cost at the time of the insurance choice in 2007, we use the predicted costs based on our prediction model estimated on the 2002-2006 data. However, this would be a lower bound on the expected cost if the increase in unemployment in 2008 at the start of the recession was anticipated. From the demand side, if aggregate unemployment risk is realized by the time that individuals choose to buy insurance, we would miss the ex-ante value of insurance against those aggregate risks by using the observed willingness-to-pay. Incorporating this ex-ante value as in Hendren [Forthcoming] would thus increase the average welfare-relevant value from extra coverage

[43] See Handel et al. [2019] for a welfare analysis that combines frictional demand and pricing inefficiencies due to adverse selection.

[44] As mentioned before, Hendren [2017] finds that adverse selection in the US context is sufficiently strong to explain the absence of private insurance markets. In general, this prediction depends on the minimum mandate in place (see footnote 3.2), which is more generous in the US than in Sweden. Note also that Hendren [2017]'s test does not account for other sources of heterogeneity that may reduce the correlation between risk and willingness-to-pay, which we show to be relevant in the Swedish context.

relative to the estimated demand curve.

7 Conclusion

Seventy five years ago, the Beveridge Report, in its attempt at increasing welfare for all, recommended a new set of revolutionary social insurance policies and insisted that these insurance policies "must be achieved by co-operation between the State and the individual. (...) The state should not stifle incentive, opportunity, responsibility; in establishing a national minimum, it should leave room and encouragement for voluntary action by each individual to provide more than that minimum for himself and his family".[45] In the context of unemployment insurance though, generous mandates have left very little room for choice and voluntary actions by individuals, under the (untested) presumption that offering choice would trigger risk-based selection. Our paper provides the first direct evidence for risk-based selection into unemployment insurance and offers new insights on how to reconcile social insurance design with individual choice in such a context of adverse selection.

Using various empirical strategies and different sources of variation, we robustly find that workers who face higher (ex-ante) unemployment risk select into more comprehensive coverage. Despite the severe adverse selection, we find that mandating all workers into comprehensive coverage is not the best policy response. The Swedish workers who choose not buy the comprehensive coverage value it below its cost, which is high due to moral hazard. This is an important result, given an ongoing policy debate in Sweden regarding a fully mandated unemployment insurance system, but also given the universal mandates of comprehensive UI that are in place around the world – the desirability of which has never been tested before. Moreover, the ubiquitous absence of private markets for UI may be precisely because of the mandated public programs in place and the adverse selection into supplemental coverage they cause. As our analysis shows, the impact of adverse selection can be mitigated by subsidizing the premia for more comprehensive coverage, which would increase the desirability of a more generous minimum mandate.

Importantly, the value of providing choice in social insurance programs does rely on individuals using this choice in their best interest. This is an assumption we have maintained throughout to focus our analysis. A rapidly growing literature documents the importance of frictions distorting households' insurance choices. This introduces another important caveat when introducing choice, closely related to the potential for adverse selection. We leave this challenging, but important issue for future research.

[45] Beveridge [1942]

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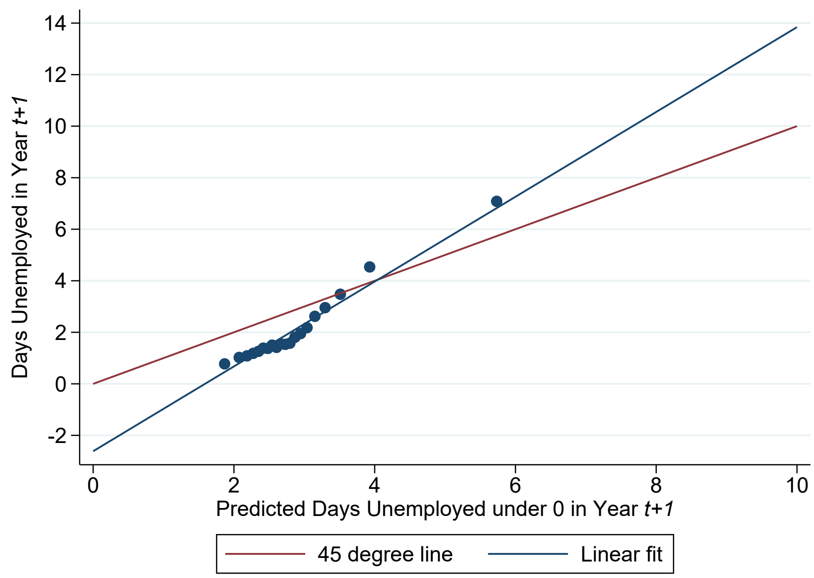
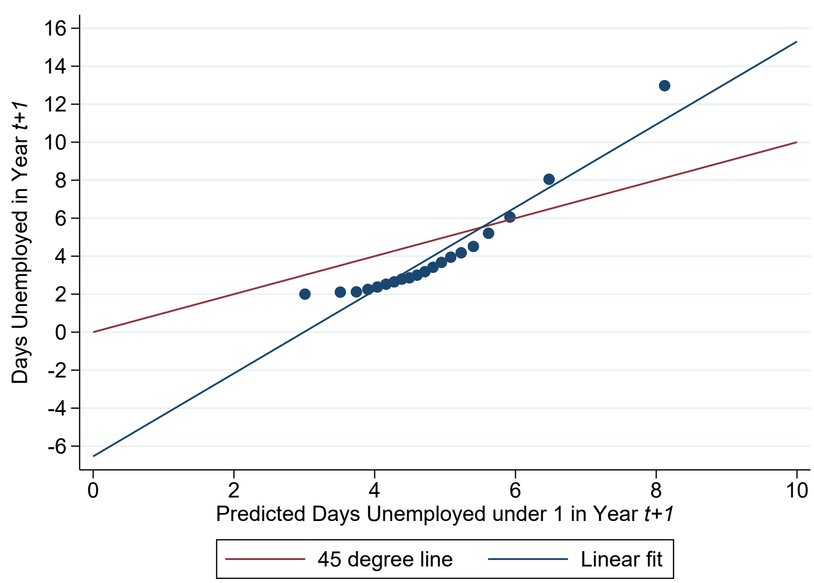
Figure 1: Predicted Unemployment Risk: Model Fit

A. Predicted vs Realized Risk Under Basic Coverage

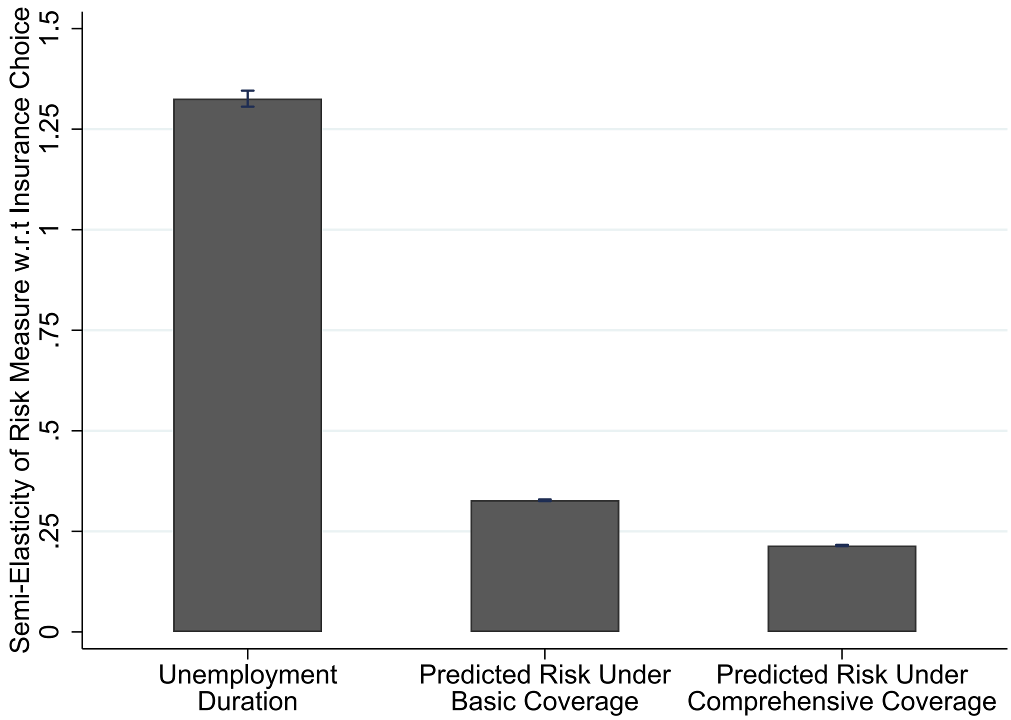


B. Predicted vs Realized Risk Under Comprehensive Coverage



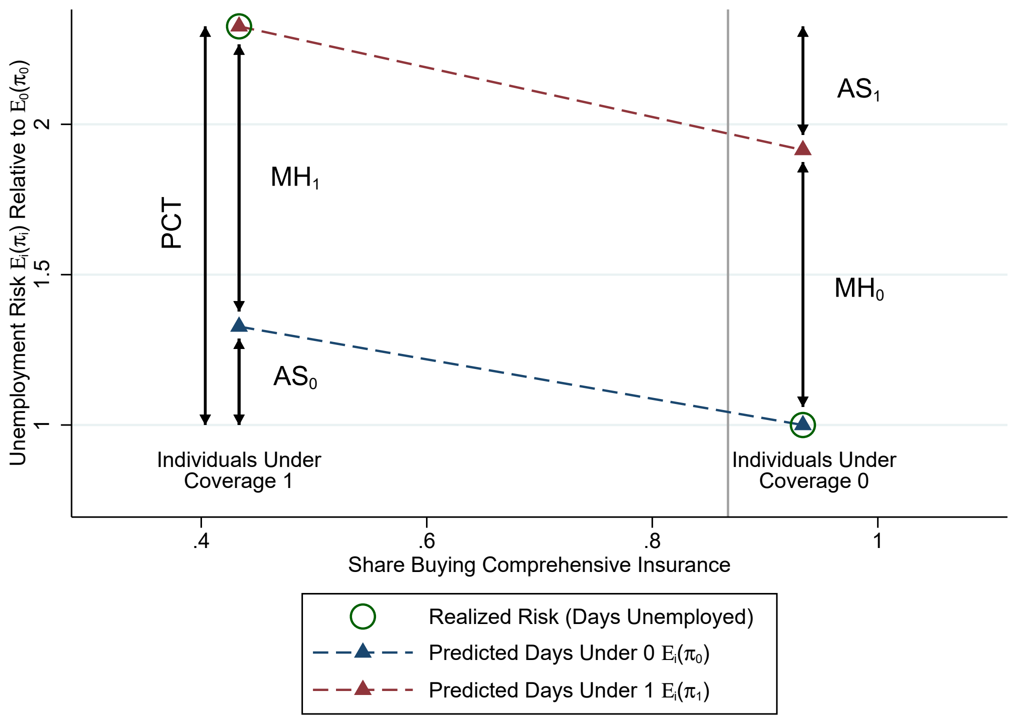
Notes: The Figure reports binscatters correlating realized unemployment risk and measures of predicted unemployment risk, from the model presented in Section 2.3. The model combines flexibly all observable sources of risk, including institutional shifters of risk such as the full history of the firm layoff notifications, and the relative tenure ranking of the individual. Model selection is based on the Lasso approach for zero-inflated poisson suggested by Banerjee et al. [2018]. To allow for moral hazard, we estimate a model of risk for individuals under the basic coverage, and a separate model of risk for individuals under the comprehensive coverage. The model predicts the number of days spent unemployed in year $t+1$ based on observable characteristics in year t. Panel A correlates predicted unemployment risk under the basic coverage ($\hat{\pi}_0$) with realized unemployment risk in $t+1$ for individuals observed under basic coverage in year t. Panel B correlates predicted unemployment risk under the comprehensive coverage ($\hat{\pi}_1$) with realized unemployment risk in $t+1$ for individuals observed under comprehensive coverage in year t.

Figure 2: Positive Correlation Tests: Results



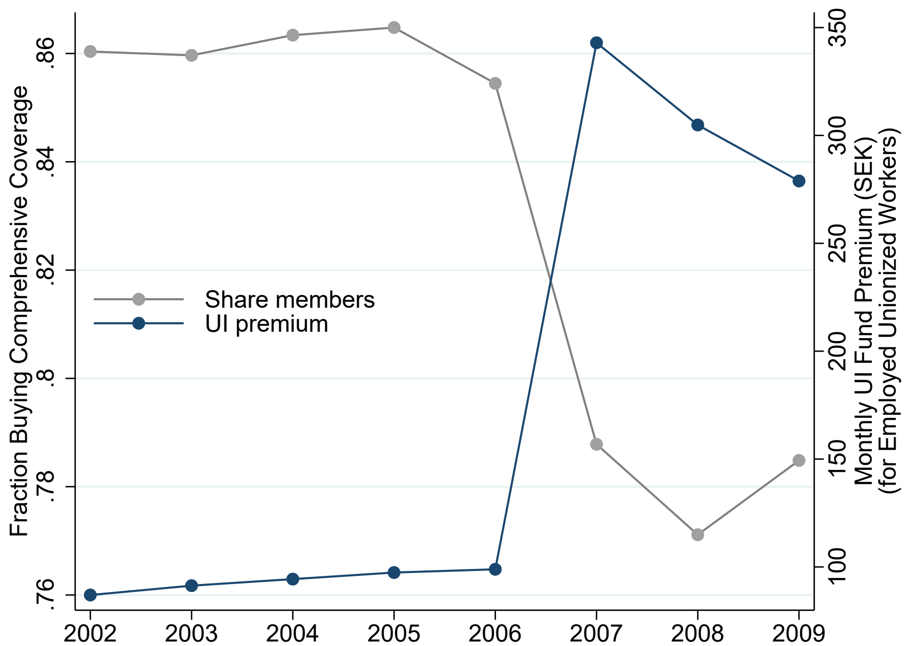
Notes: The Figure reports estimates of positive correlation tests following specification (12) estimated over the period 2002-2006 for three different risk outcomes: total duration spent unemployed in $t+1$ (π), predicted risk under basic coverage ($\hat{\pi}_0$), and predicted risk under comprehensive coverage ($\hat{\pi}_1$). Specification (12) controls for year fixed effects and for the limited set of characteristics that affect the unemployment insurance coverage available to individuals: a dummy for whether individuals meet the work eligibility requirement, a dummy for union membership, and earnings level. For each outcome, the chart displays the semi-elasticity of the risk outcomes with respect to insurance choices in t defined in (13). See text for details.

Figure 3: Decomposition of PCT between AS and MH



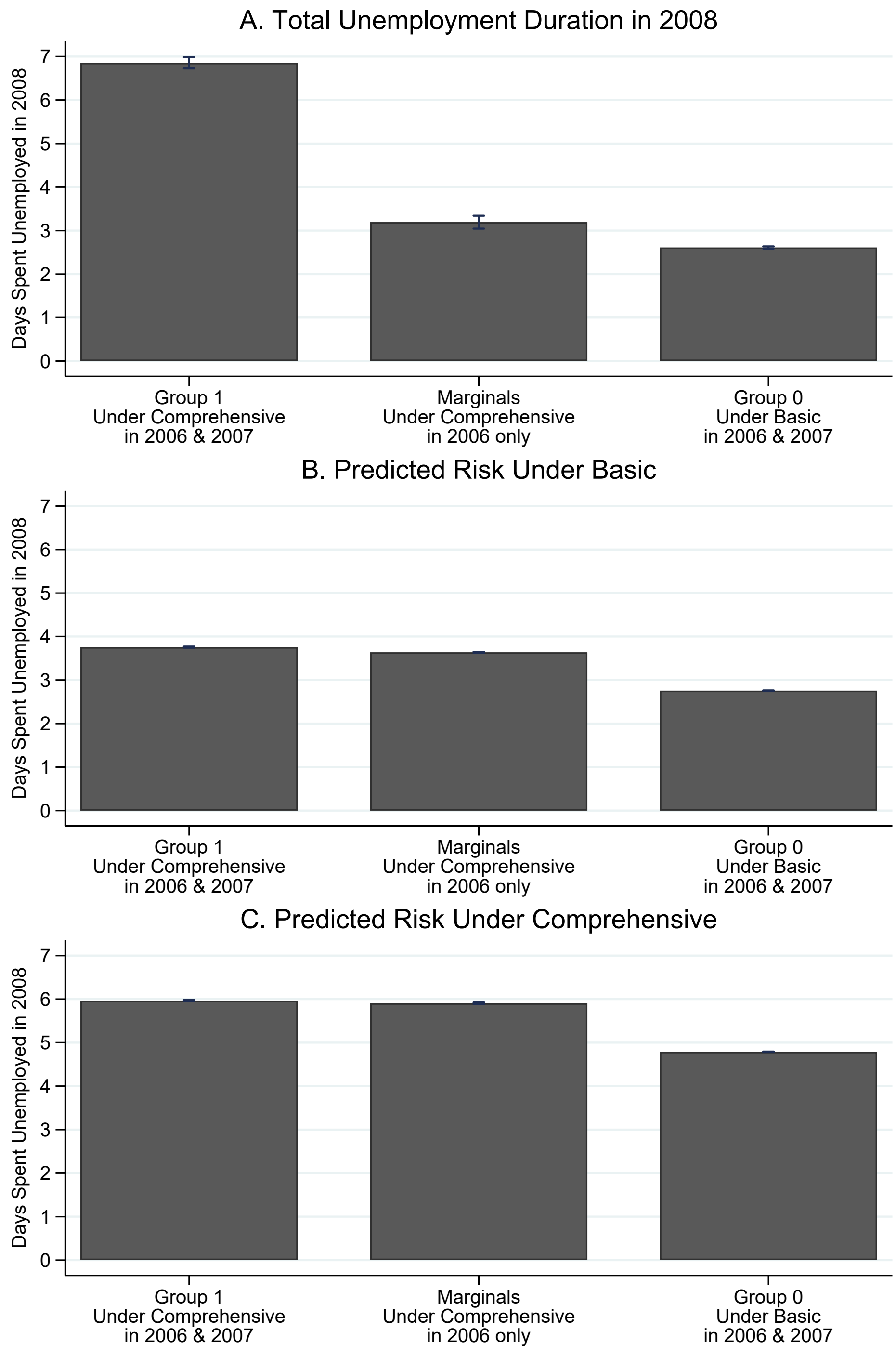
Notes: The Figure uses the PCT estimates of Figure 2 to offer an implementation of the decomposition of the PCT following the model of Figure F.1. Green dots represent the observed average realized risk, measured by the number of days spent unemployed in year $t+1$. The difference in realized risk $E_1(\pi_1) - E_0(\pi_0)$ is obtained from the semi-elasticity (13) and corresponds to our baseline positive correlation test-statistic reported in the first bar of Figure 2. All risk measures are conditional on the vector of characteristics Z, and normalized to the average risk under basic coverage of individuals observed under basic coverage $E_0(\pi_0)$. Blue triangles represent the average predicted risk under basic coverage for individuals selecting basic coverage, $E_0(\hat{\pi}_0)$, and for individuals selecting comprehensive coverage, $E_1(\hat{\pi}_0)$. The difference in predicted risk under basic coverage for the two groups, $E_1(\hat{\pi}_0) - E_0(\hat{\pi}_0)$ is obtained from the semi-elasticity (13) using the predicted risk $\hat{\pi}_0$ as an outcome. The red triangles replicate the same exercise for predicted risk under under comprehensive coverage.

Figure 4: Price Variation: Evolution of Premia **p** and of the Fraction of Workers Buying the Comprehensive Coverage Around the 2007 Reform



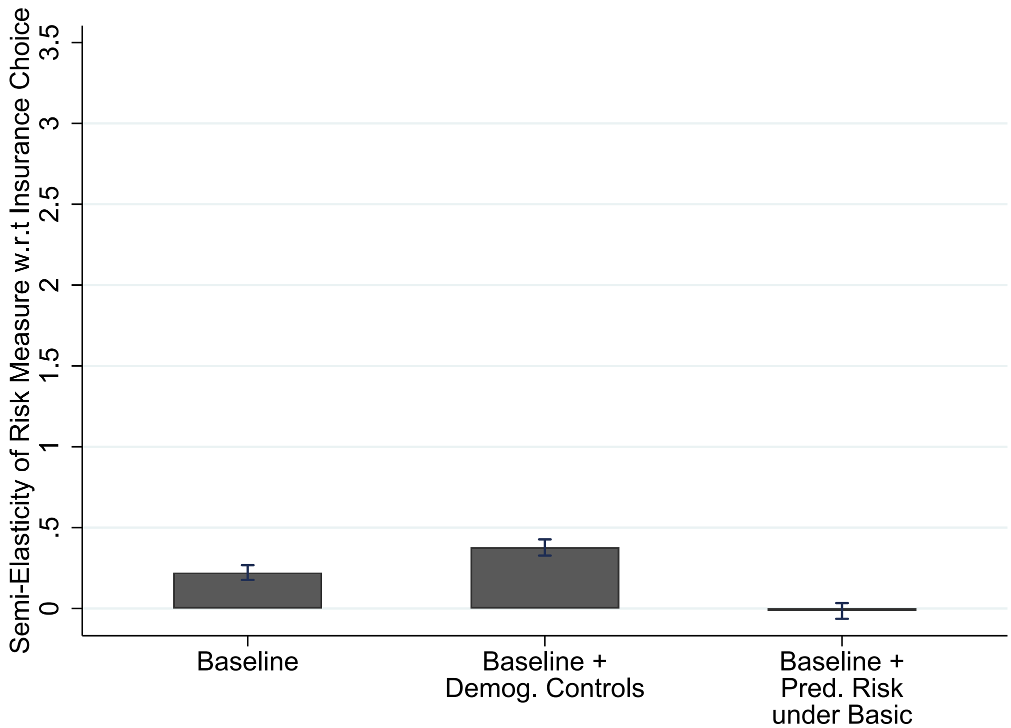
Notes: The Figure reports the evolution of monthly premium for the supplemental UI coverage over time. As explained in Section 2.1, there are no sources of premium differentiation up to 2008, apart from small rebates for union members and for unemployed individuals. Here, we report the value of the premium for employed union members. The Figure shows a large and sudden increase in the premia paid for the supplemental coverage in 2007. This increase followed the surprise ousting of the Social Democrats from government after the September 2006 general election. Note that from July 2008 on, premia started to be differentiated across UI funds. For 2008 and 2009 we therefore report the average monthly premium among unemployed union members across all UI funds. The Figure also shows the evolution of the take-up of the comprehensive UI coverage, measured as the sum of all individuals buying the comprehensive coverage divided by the total number of individuals aged 25 to 55 meeting the eligibility criteria for receiving UI benefits.

Figure 5: Price Variation: Unemployment Risk by Willingness-To-Pay



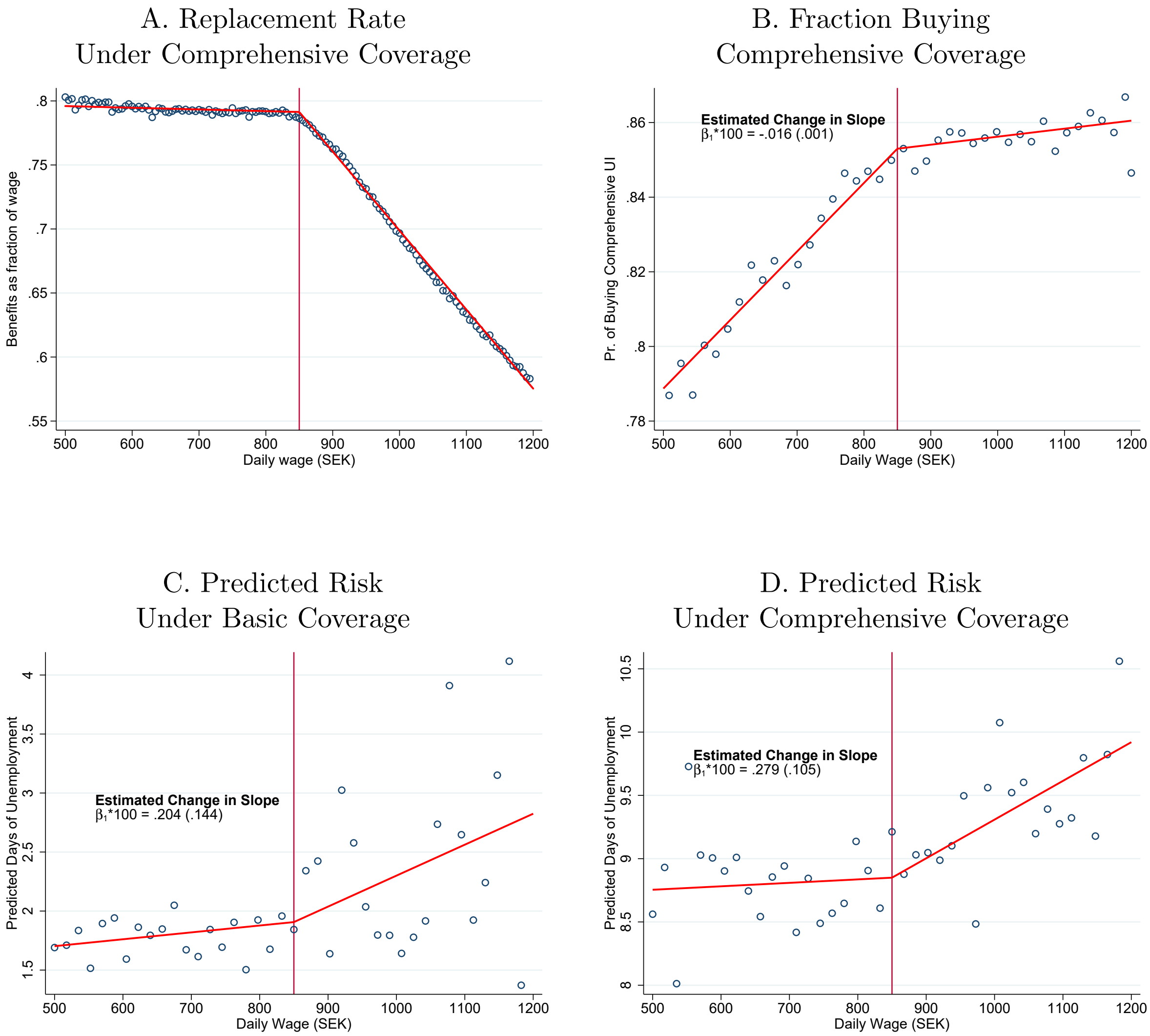
Notes: The Figure reports average risk for three groups of individuals defined by descending order of willingness-to-pay. Group 1 individuals are in the comprehensive coverage both in 2006 and 2007: they have the highest valuation of the supplemental coverage ($\mathbf{u} > 0$). Marginals $M(\mathbf{p})$ were buying the comprehensive coverage in 2006 but switch out in 2007 when premia $\mathbf{p}$ increase: they are close to indifferent between the two coverages at current prices ($\mathbf{u} \approx 0$). Individuals from group 0 were neither buying the comprehensive coverage in 2006 nor in 2007, and have the lowest willingness-to-pay for the supplemental coverage ($\mathbf{u} < 0$). For each panel, we report $E(\pi | Z = \bar{Z}_0, D = j)$, the average risk outcome π of each group $j = 1, M, 0$ estimated at the average value of Z for group 0. The vector Z are characteristics affecting contract differentiation. Panel A reports the average number of days spent unemployed in 2008 for each group. Panel B plots the average predicted risk under basic coverage, and panel C the average predicted risk under comprehensive coverage. See text for details.

Figure 6: Price Variation: Test of Residual Private Information



Notes: The Figure uses the 2007 price reform to estimate how much adverse selection is left when controlling for observable characteristics. The graph first reports the baseline semi-elasticity $\text{Semi}_{M(\mathbf{p})}^{Baseline}$ of realized risk for the marginals M relative to the individuals always in the basic coverage. The second bar shows the same semi-elasticity where we include in specification (16) a set of dummies for age, gender, marital status, education (four categories), industry (1-digit code) and wealth level (quartiles). Interestingly, the semi-elasticity increases compared to our baseline when including these controls. This suggests these characteristics provide advantageous selection on average, such that if contracts were differentiated along these observable dimensions, adverse selection into comprehensive coverage would actually be more severe. The third bar reports the semi-elasticity $\text{Semi}_{M(\mathbf{p})}^{Residual}$ where we condition on predictable risk by including in specification (16) twenty dummies for the predicted unemployment risk score under basic coverage in both the inflated and count part of the model. Results indicate that there is no significant residual correlation left between realized risk and willingness-to-pay when we control for predicted unemployment risk. See text for details.

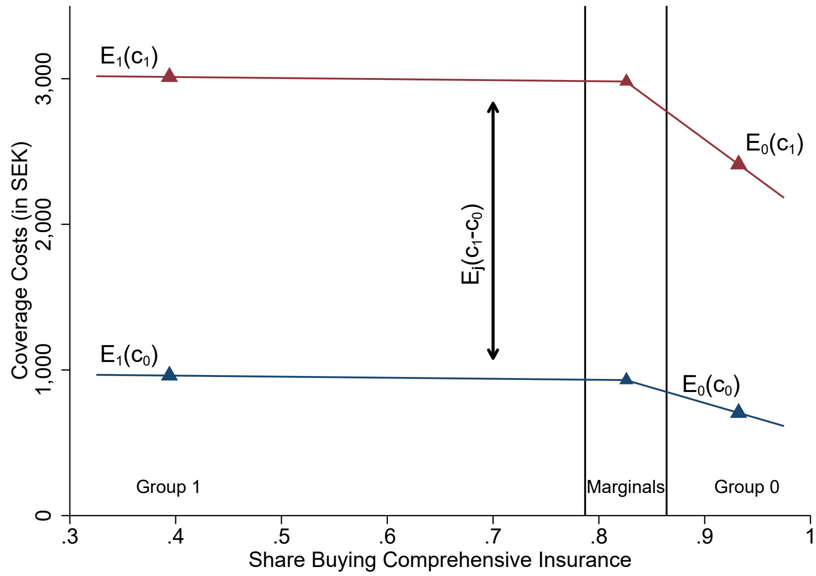
Figure 7: Benefit Variation: Regression Kink Design Analysis of Demand Responses and Risk Based Selection



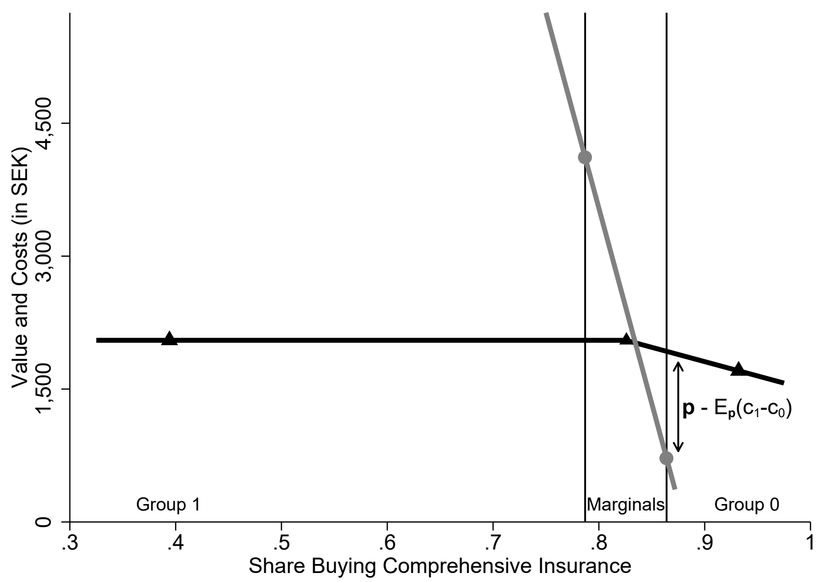
Notes: The Figure presents the regression kink design based on the kinked schedule of benefits b_1 in the comprehensive coverage. The sample consists of all individuals from our baseline sample who become unemployed in year $t \in [2003; 2007]$ and were employed in year $t-1$. Panel A shows the relationship between the daily wage prior to unemployment and the replacement rate, defined as average daily benefit received during unemployment from the IAF data divided by the daily wage. The panel shows that the relationship exhibits a sharp kink at $w = 850SEK$, the level of the cap in b_1. Panel B plots the average fraction of individuals who bought the comprehensive coverage in year t by bins of the daily wage in year t. The data is residualized on the vectors of characteristics Z which affect the contract space and on the vector X which includes age, gender, level of education, region, family type and industry. We report on the graph the linear fit and the estimate β_1 corresponding to specification (18) estimated using a bandwidth of 350SEK. Panels C and D produce RKD graphs similar to panel B, using as outcomes predicted risk under basic coverage in year t and predicted risk under comprehensive coverage in year t respectively.

Figure 8: Welfare Analysis

A. Cost Curves for Comprehensive and Basic Plan



B. Demand vs. Cost of Supplemental Coverage



Notes: The figure shows the demand and cost curves underlying the welfare implementation in Section 6. Panel A plots the cost curves for the comprehensive and basic plan. For the basic (comprehensive) cost curve, we convert the predicted days spent unemployed under basic (comprehensive) coverage in Figure (5) into the expected cost of providing basic (comprehensive) coverage for individuals from group 1, marginals and individuals from group 0 defined using the 2007 price change. Following Figure (F.2), we locate the estimated average costs for the three groups at the midpoint of the ranges of the corresponding valuation quantiles and then use a piece-wise linear interpolation to construct the cost curves. Panel B compares the willingness-to-pay and cost $\mathbf{c}$ of providing supplemental coverage for different workers ranked based on their valuation $\mathbf{u}$. The black triangles correspond to the difference between the cost of providing the comprehensive plan and the basic plan (plotted separately in Panel A) for each of the three groups. The supplemental cost curve is then constructed using a piece-wise linear interpolation. The linear demand curve is an extrapolation of the share of individuals taking up the comprehensive plan at the pre- and post-reform prices (grey circles). The vertical distance between the demand and cost curves in Panel B denotes the fiscal externality per marginal worker $FE^{AS}_{\mathbf{p}} = [p_1 - p_0] - \left[E_{M(\mathbf{p})}(\pi_1)\, b_1 - E_{M(\mathbf{p})}(\pi_0)\, b_0\right]$ for different prices $\mathbf{p}$.

Table 1: Summary Statistics

	Mean	P10	P50	P90
	I. Unemployment			
Displacement probability	3.05%	-	-	-
Displacement probability (exc. quits)	2.85%	-	-	-
Unemployment probability	3.65%	-	-	-
Days unemployed (unconditional)	5.28	0	0	0
Unemployment spell duration (days) (if > 0)	147.39	22	91	305
Predicted days unemployed under Basic	3.57	2.14	2.72	4.02
Predicted days unemployed under Comprehensive	5.83	3.63	4.77	7.17
Fraction receiving layoff notification	.05	-	-	-
Fraction switching firms	.09	-	-	-
	II. Union and UI Fund Membership			
Union membership	.75	-	-	-
UI fund membership (D)	.86	-	-	-
Fraction switching from coverage 0 to 1	.02	-	-	-
Fraction switching from coverage 1 to 0	.01	-	-	-
	III. Demographics			
Age	40.88	28	40	52
Years of education	12.86	11	12	16
Fraction men	.52	-	-	-
Fraction married	.44	-	-	-
	IV. Income and Wealth SEK 2003(K)			
Gross earnings	249	100.3	231.2	388.8
Net wealth	385.9	-160.8	90.9	1127.7
Bank holdings	49.3	0	0	122.1
N	17,761,796			

Notes: The Table provides summary statistics for our baseline sample over the period 2002 to 2006. The sample comprises individuals who are between 18 and 60 years of age and have been working for at least 6 months. It contains 17.8 million worker $\times$ year observations. Unemployment is defined as a spell of non-employment, following an involuntary job loss, and during which an individual has zero earnings, receives unemployment benefits and reports searching for a full time job. We define displacement as an involuntary job loss, due to a layoff or a quit following a 'valid reason'. Voluntary quits are not included in our measures of displacement and unemployment. The probability of displacement is the probability to be displaced in year $t+1$ conditional on working in year t. The unemployment probability is the probability to be unemployed in year $t+1$ unconditional on employment status in year t. The fraction of workers receiving layoff notification comes from the layoff-notification register (VARSEL) and is defined as the fraction of workers that are employed in an establishment emitting a layoff notification in year t. The employer-employee matched data (RAMS) registers all existing labor contracts on a monthly basis. We define a "firm switch" as moving from having a labor contract with firm j to having a contract with firm k, without any recorded non-employment spell between these two contracts. UI fund membership information comes from tax data for the period 2002 to 2006, during which premia were eligible for a 40% tax credit. The dummy variable D for buying the comprehensive coverage in year t is defined as reporting any positive amount of premia paid in year t. All earnings, income and asset level measures are from wealth and income registers, and are yearly measures in constant k2003SEK. All assets are aggregated at the household level and estimated at their market value. 1SEK2003 $\approx$ 0.11 USD2003

Appendix A Observable Risk & Predicted Risk Model: Further Material

In this appendix:

(i) we present the relationship in the Swedish labor market, between a series of observable risk shifters, that are credibly exogenous to individuals' own actions, and individuals' unemployment risk.

(ii) we then present a model of predicted unemployment risk that combines all risk shifters together.

A.1 Observable Risk Shifters in the Swedish Labor Market

The institutional context and the richness of the Swedish registry data allows us to observe determinants of unemployment risks that are arguably beyond the control of individuals. We present here three such observable sources of risk variation and show how they correlate with individuals' realized unemployment risk.

Average Firm Layoff Rates The first observable source of variation in unemployment risk stems from firm level risk. Firm level risk can vary cross-sectionally, due to permanent differences in turnover across firms, or over time, due to firms experiencing temporary shocks. We focus, to start with, on the permanent component of firm level risk, and explore how this permanent component correlates with an individual's displacement probability.

For each individual i working in firm j, we define average firm displacement risk $\bar{\pi}_{-i,j}$ as the average probability of displacement of all other workers within the firm excluding individual i over all years where the firm is observed active between 1990 and 2015.[46] We then plot, in Figure A.1 panel A, our measure of average firm risk $\bar{\pi}_{-i,j}$ in 20 bins of equal population size, against $\pi_{i,j}$, the individual probability of displacement in $t+1$, for all individuals ever employed during the period 2002-2007. The figure shows first that there is significant heterogeneity in firms' average separation rates. Second, the figure provides clear evidence that individuals' unemployment risk is very strongly correlated with average firm level risk.

Layoff Notifications The second observable source of variation in unemployment risk stems from variation in firm level risk over time. We leverage the fact that under Sweden's employment-

[46] For this purpose we match the employer-employee registry (RAMS) from 1985 to 2015 with the Public Employment Service (PES) registry for all years 1990-2015.

protection law, firms subject to a shock and intending to displace 5 or more workers simultaneously must notify the Public Employment Service in advance. Once a notification is emitted, employers need to come up with the list and dates for the intended layoffs. These layoffs may happen up to 2 years after the original notification has been sent. The layoff notifications provide a source of observable variation in firm displacement risk. In Figure A.2 we report the evolution of the displacement probability of workers around the first layoff notification event in the history of the firm. We define event year n as the year to/since the firm emits its first layoff notification and follow an event study approach around that event. Our sample is the panel of workers who are employed in the firm at the date this layoff notification is emitted to the PES. The graph shows that a layoff notification is indeed associated with a sudden and large increase in the displacement probability of workers. Immediately following the layoff notification, the displacement rate of workers jumps by 6 percentage points compared to pre-notification levels, and remains high for about two years, before decreasing and converging back to pre-notification levels.

Because the panel of workers is selected based on being employed in the firm in year $n = 0$, one may worry that this surge in displacement rates is mechanical, as displacement can only increase after year 0 conditional on all workers being employed in year 0. To mitigate this concern, we follow a matching strategy and create a control panel of workers selected along the same procedure as the original panel. We use nearest-neighbor matching to select a set of firms that are similar, along a set of observable characteristics, to the firms emitting a layoff notification, but never emit a layoff notification.[47] We allocate to the matched firm in the control group a placebo event date equal to the layoff notification date of her nearest-neighbor in the treated group of firms. We then select workers that are in the control firm at the time of the placebo event date to create our matched control panel. Results in Figure A.2 show that, pre-event, the displacement risk is very similar in the control and treated groups, and that it evolves smoothly in the control firms around the event.

This evidence suggests that layoff notifications are a significant shifter of individuals' unemployment risk, as they immediately double the baseline displacement probability of workers.

Relative Tenure The third source of observable (and credibly exogenous) risk variation is at the individual level and stems from the strict enforcement of the Last-In-First-Out (LIFO) principle. When a firm wants to downsize within an establishment, the legal system prescribes that displacement occurs by descending order of tenure within the establishment. In practice, workers are divided into groups, defined by collective bargaining agreements in the establishment, and then a tenure ranking within each group is constructed. The tenure ranking of an individual within her establishment and collective bargaining agreement (CBA) group directly determines her probability to be separated. A limitation of our data is that workers' collective bargaining agreements are not directly observed. Instead, we use detailed occupation codes as proxies for the CBAs, which is a good approximation as most CBAs are done at the occupation level.

Figure A.1 panel B plots the probability of being displaced in $t + 1$ among individuals working

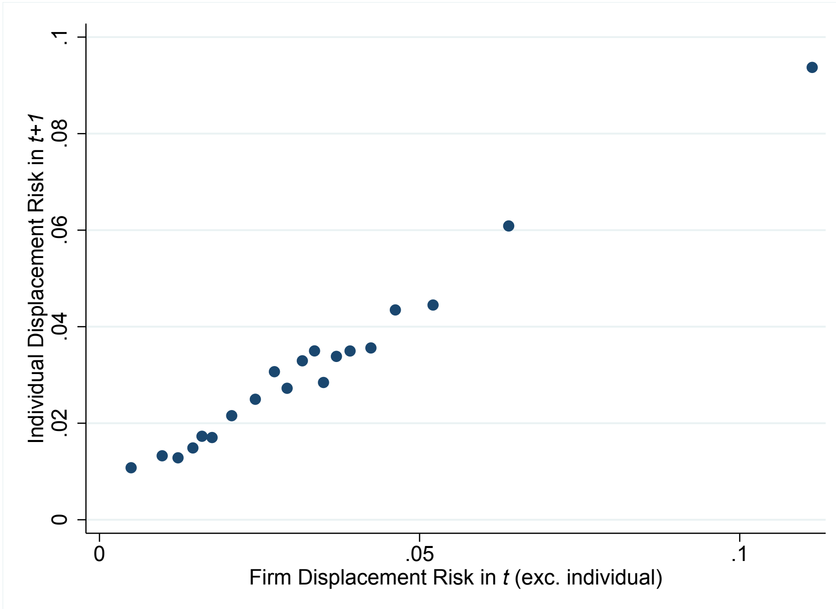
[47] The covariates used for matching are the number of employees, the 4 digit sector codes of the firm, the average earnings and average years of education of workers in the firm.

in firms that emit a layoff notification in $t+1$, as a function of relative tenure ranking within establishment and occupation. The Figure provides clear evidence of a strong negative correlation between relative tenure ranking and individuals' displacement probability. Individuals within the lowest 10 percent of tenure rankings have a probability of being displaced in $t+1$ larger than .1; this probability declines steadily as tenure ranking increases, and then stays below .02 for individuals in the highest 50 percent tenure rankings.

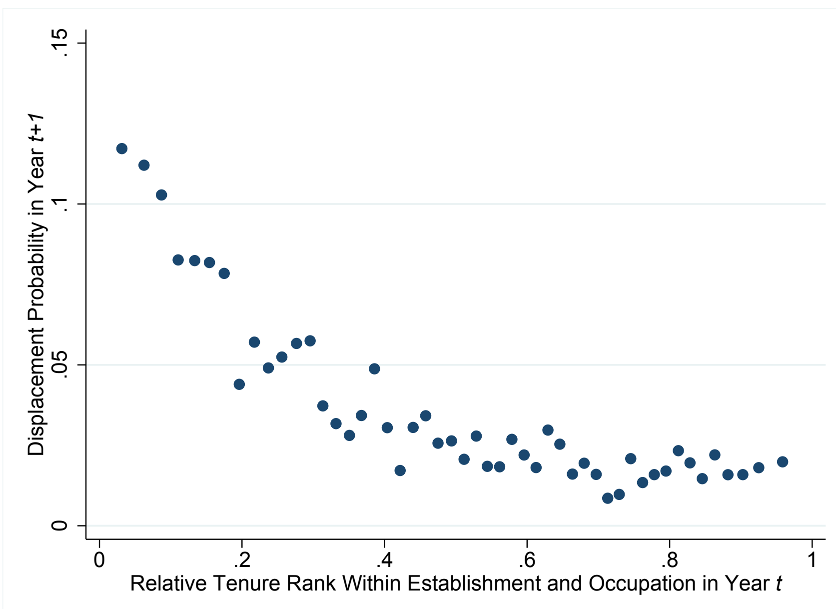
For firms with multiple establishments, one layoff notification needs to be sent for each establishment intending to layoff workers and the LIFO principle applies at the level of the establishment. While the institutionalization is specific to Sweden, the LIFO principle is used for determining redundancy in many countries (e.g., Netherlands, Poland, UK, etc).

Figure A.1: Risk Shifters: Firm Displacement Risk & Last-In-First-Out Principle

A. Firm Displacement Risk vs Individual Displacement Probability in $t+1$

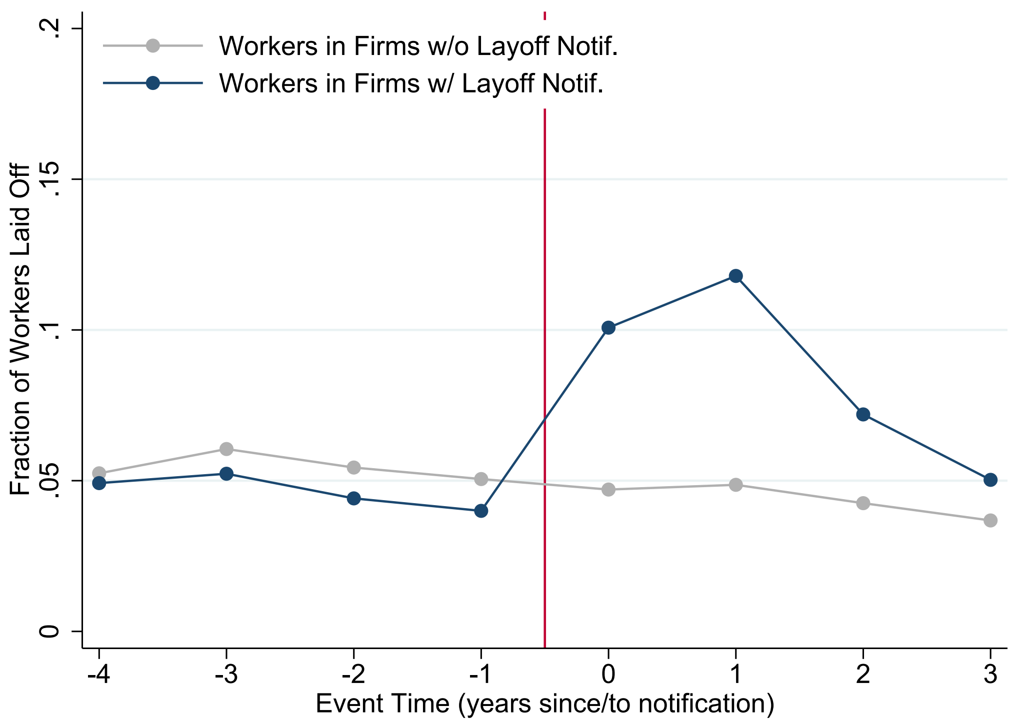


B. Relative Tenure Ranking in Year t vs Displacement Probability in Year $t+1$



Notes: The Figure provides evidence of the role of firm level risk and of the Last-In-First-Out (LIFO) principle in creating variation in individuals' unemployment risks. In panel A, we provide evidence of the role of firm layoff risk as a shifter of individuals' own displacement probability. For each individual i working in firm j, we define average firm displacement risk as the average probability of displacement of all other workers within the firm excluding individual i, $\bar{\pi}_{-i,j}$ over all years where the firm is observed active in our sample years. We then plot the average firm displacement risk in 20 bins of equal population size, against the individual probability of displacement in $t+1$. The Figure shows that there is significant heterogeneity in firms' separation rates, and that individuals' unemployment risk is very strongly correlated with firm level risk. Panel B plots the probability of being displaced in $t+1$ among individuals working in firms that emit a layoff notification in $t+1$, as a function of relative tenure ranking within establishment and occupation in year t. See Section 2.1 for institutional details. The Figure provides clear evidence of a strong negative correlation between relative tenure ranking and individuals' displacement probability.

Figure A.2: Layoff Notification and Displacement Risk



Notes: The Figure reports estimates of the evolution of the displacement probability of workers around the first layoff notification event in the history of the establishment. We define event year $n = 0$ as the year in which an establishment emits its first layoff notification, and focus on the panel of workers who are employed in the establishment at the date this layoff notification is emitted to the PES. The graph shows that a layoff notification is indeed associated with a sudden and large increase in the displacement risk of workers. Because the panel of workers is selected based on being employed in the firm in year $n = 0$, one may worry that this surge in displacement rates is mechanical, as displacement can only increase after year 0 conditional on all workers being employed in year 0. To mitigate this concern, we follow a matching strategy and create a control panel of workers selected along the same procedure as the original panel. We use nearest-neighbor matching to select a set of firms that are similar, along a set of observable characteristics, to the firms emitting a layoff notification, but never emit a layoff notification. We allocate to the matched firm in the control group a placebo event date equal to the layoff notification date of her nearest-neighbor in the treated group of firms. We then select workers that are in the control firm at the time of the placebo event date to create our matched control panel.

A.2 Predicting Risk Using Observable Risk Shifters: a Zero-Inflated-Poisson Model of Unemployment Risk

We now present the model we use to compute the best predictor of future unemployment risk given all currently observed individual characteristics. To do so, we leverage the rich set of observables available in the Swedish registry data, and the various institutional features of the Swedish labor market.

Setup The measure of unemployment risk π that we model is the number of days an individual is expected of spending unemployed in $t+1$. This is the relevant measure of risk to the UI system given insurance choices made in year t.[48]

The distribution of days spent unemployed is defined only over non-negative integers, and exhibits a significant mass at zero. To account for these facts, we model π using a zero-inflated Poisson model. The expected number of days unemployed conditional on a vector of characteristics X therefore takes the following form:

$$E(\pi|X) = (1 - f(0|X^I)) \exp(X^{C'} \beta^C)$$

For the zero-inflated part of the process, we parametrize the probability $f(0)$ using a logit: $f(0|X^I) = \exp(X^{I'}\beta^I)/(1+\exp(X^{I'}\beta^I))$. We will allow for the set of risk predictors X^I and X^C entering respectively the inflated part and the count part, to differ.

To account for moral hazard, we allow the risk of individuals with similar characteristics X to differ if they are observed under the basic coverage or under the comprehensive coverage. To this purpose, we estimate separately two models of predicted risk. The first model is the predicted risk given X under the basic coverage $\hat{\pi}_0 = E(\pi_0|X)$. This model is estimated on individuals who are observed under the basic coverage in t. The second model is the predicted risk given X under the comprehensive coverage $\hat{\pi}_1 = E(\pi_1|X)$, which we estimate on individuals who are observed under the comprehensive coverage in t.

Lasso Penalization In terms of model selection, we discipline the choice of the many potential regressors by using the adaptive Lasso procedure for a zero-inflated Poisson (ZIP) model proposed by Banerjee et al. [2018]. The ZIP Log Likelihood function with LASSO penalty works in the following way.

Let $X^C = \{x_1^C, ..., x_K^C\}$ be the set of K regressors associated with predicting the number of days unemployed, conditional on some unemployment, according to a Poisson distribution. The corresponding coefficients are: $\{\beta_1^C, ..., \beta_K^C\}$. Let $X^I = \{x_1^I, ..., x_J^I\}$ be the set of J regressors associated with predicting some unemployment, according to a logistic distribution. The corresponding

[48] If an individual has bought the comprehensive coverage throughout year t, then the days she spends unemployed in year $t+1$ will be covered by the comprehensive benefits. In that sense, the relevant risk to determine the cost of providing the comprehensive coverage to an individual buying that coverage in year t is the expected number of days she will spend unemployed in year $t+1$.

coefficients are: $\{\beta^I_1, ..., \beta^I_J\}$. The number of days (integers) spent unemployed by individual i is π_i. Then we can write down the ZIP Log Likelihood function with LASSO penalty as follows:

$$L = L_1 + L_2 - L_3 - P_C - P_I$$

Where each of the L_1, L_2, L_3 components are defined as follows:

$$L_1 = \sum_{i:\pi_i=0}^{n} \log[\exp(X_i'^I \beta^I) + \exp(-\exp(X_i'^C \beta^C))]$$

$$L_2 = \sum_{i:\pi_i>0}^{n} \{\pi_i X_i'^C \beta^C + \exp(-X_i'^C \beta^C) - \ln(\pi_i!)\}$$

$$L_3 = \sum_{i=1}^{n} \log[1 + \exp(X_i'^I \beta^I)]$$

And the P_C, P_I components are defined as follows,

$$P_C = \lambda_C \sum_{k=1}^{K} |\beta^C_k|$$

$$P_I = \lambda_I \sum_{j=1}^{J} |\beta^I_j|$$

Estimation We can then estimate the model for various levels of penalization for the count part λ_C and the inflated part of the model λ_I. In practice, we draw 50 pairs of λ_C and λ_I. The largest pair is chosen so that all variables except the constant are set to zero. This corresponds to the largest penalization.

We then randomly select a subset of observations from our sample to obtain a training sample, the rest of the observation is considered our test sample. We estimate the model on the training sample for all 50 pairs of λ_C and λ_I. We then compute the MSE on the test sample for all 50 models, and select the lambda pair associated with the smallest MSE on the test sample.

Predictors The regressors we allow to initially enter the model are individual log earnings, family type, nine age bins, gender, twelve dummies for education level, year fixed effects, region fixed effects, industry fixed effects, dummies for the past layoff history of the individual, dummies for the layoff notification history of the firm, the leave-out mean of firm layoff risk, union membership, tenure rank, interactions between tenure ranking and firm layoff risk and interactions between tenure ranking and layoff notification history of the firm. We allow all these predictors to enter in both the count and inflated part.

When varying the level of penalization in the model, starting from the highest penalization, we can see what variables are the strongest predictors of the inflate and count part of unemployment risk. For the inflate component, the first variables to become significant are firm layoff risk, layoff notification dummies, and relative rank. This confirms the important role played by the institutional features of the Swedish labor market in determining unemployment risk. For the count component, the first variables to become significant (by order) are those associated with age, gender, education level, income, regions and years

The results show that the optimal penalization factors λ associated with the count component are smaller while those associated with the zero component are higher, thus penalizing the inclusion of variables in the latter more. As a result, in our preferred model of predicted risk, in the zero component, a large share of variables have a coefficient set to zero.

Model fit As explained in section 2.3, we first assess the quality of the model fit in Figure 1 by plotting bin scatters of the relationship between predicted risk under basic (resp. comprehensive) coverage and actual realized risk for individuals under basic (resp. comprehensive) coverage. In both panels, the relationship is close to the 45 degree line indicating that the model does a good job at predicting the average realization of unemployment risk. In Table A.1 below, we provide further summary statistics on the distribution of predicted risk according to our model. In Panel A, we focus on individuals observed under the basic coverage, and compare the distribution of their realized risk π_0 to the distribution of their predicted risk under basic coverage $\hat{\pi}_0$. The average risk predicted by the model (2.95) is very close to the average realized risk (2.83). In Panel B, we do a similar exercise, focusing on individuals observed under the comprehensive coverage. We compare the distribution of their realized risk π_1 to the distribution of their predicted risk under basic coverage $\hat{\pi}_1$. We find again that the average risk predicted by the model (5.90) is very close to the average realized risk (5.65). In both panels, we find that there is much less dispersion in predicted risk than in realized risk. The standard deviation of predicted risk is roughly six times smaller than that of realized risk. This confirms that there still remains a significant dimension of idiosyncratic unemployment risk beyond what can be predicted by even our very rich set of observables.

Table A.1: Distribution of Realized and Predicted Risk for Individuals under Basic and under Comprehensive Coverage

Panel A. Predictable Risk Under Basic

	π_0	$\hat{\pi}_0$
P25	0	2
P50	0	3
P75	0	3
P90	0	4
P99	107	13
P99.9	346	32
Mean	2.84	2.95
s.d.	22.54	2.47
N	2,2296,727	2,232,136

Panel B. Predictable Risk Under Comprehensive

	π_1	$\hat{\pi}_1$
P25	0	4
P50	0	5
P75	0	6
P90	0	7
P99	184	30
P99.9	365	80
Mean	5.65	5.91
s.d.	33.17	5.86
N	15,003,779	14,879,543

Notes: The table reports moments of the distribution of predicted risk and realized risk from our sample of workers for years 2002 to 2006. Panel A focuses on individuals who are observed under the basic coverage in t. The first column reports moments of the distribution of their realized risk π_0 while the second column reports moments of the distribution of our measure of predicted risk under basic coverage $\hat{\pi}_0$. Panel B focuses on individuals who are observed under the comprehensive coverage in t. The first column reports moments of the distribution of their realized risk π_1 while the second column reports moments of the distribution of our measure of predicted risk under comprehensive coverage $\hat{\pi}_1$.

Appendix B Positive Correlation Tests: Further Results

In this appendix we present further evidence regarding the positive correlation between unemployment risk and UI choices:

(i) we present positive correlation tests using alternative risk outcomes.

(ii) we present robustness analysis for the PCT using alternative specifications and non-parametric approaches

Table B.3 provides the summary statistics for our main sample broken down by UI coverage.

B.1 Positive Correlation Tests: Alternative Risk Outcomes

We start by showing that the strong correlation between realized unemployment risk and UI choices documented in Figure 2 extends to using alternative measures of realized unemployment risk.

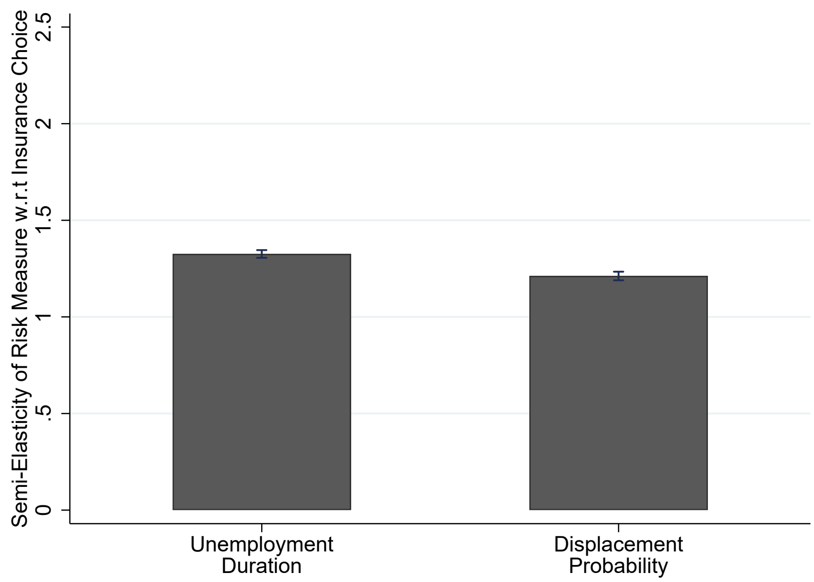
Displacement Probability We start by investigating the robustness of our PCT results to using the probability of displacement in $t+1$ as our measure of risk π. To control for observables Z, we model the probability of displacement as a probit:

$$E(\pi|Z) = \Phi(Z'\beta + \alpha \cdot \mathbb{1}[D=1]) \tag{20}$$

where $\Phi(.)$ is the standard normal c.d.f. The second bar of Figure B.1 reports the semi elasticity defined in (13) estimated from this model. The first bar of the figure reports our estimate of the PCT for our baseline measure of risk, that is the number of days spent unemployed in $t+1$. The graph confirms the presence of a strong positive correlation between UI choices and unemployment risk: Individuals who buy the comprehensive coverage in t are 125% more likely to be displaced in $t+1$ than individuals who do not buy the comprehensive coverage.

We note that different measures of unemployment risks are subject to different types of moral hazard. Comparing the magnitude of the correlations across the different realized risk outcomes already sheds light on some margins of moral hazard. A large body of literature has for example documented that higher unemployment benefits increase the duration of unemployment spells conditional on becoming unemployed (see Schmieder and von Wachter [2016] for a recent review). Such moral hazard conditional on displacement will increase the correlation between unemployment duration in $t+1$ and insurance coverage in t (first bar in Figure B.1). The probability of displacement, while immune to moral hazard once displaced, is potentially affected by moral hazard "on the job" (second bar in Figure B.1). An example of this would be collusion between employers and employees to qualify actual voluntary quits as "quits following a valid reason", which are eligible for unemployment benefits.

Figure B.1: Positive Correlation Tests - Displacement Risk vs Total Unemployment Risk



Notes: The first bar of the figure reports our estimate of the PCT for our baseline measure of risk, that is the number of days spent unemployed in $t+1$. The second bar reports the PCT for the displacement risk in $t+1$. This estimate is the semi elasticity defined in (13) estimated from probit specification (31).

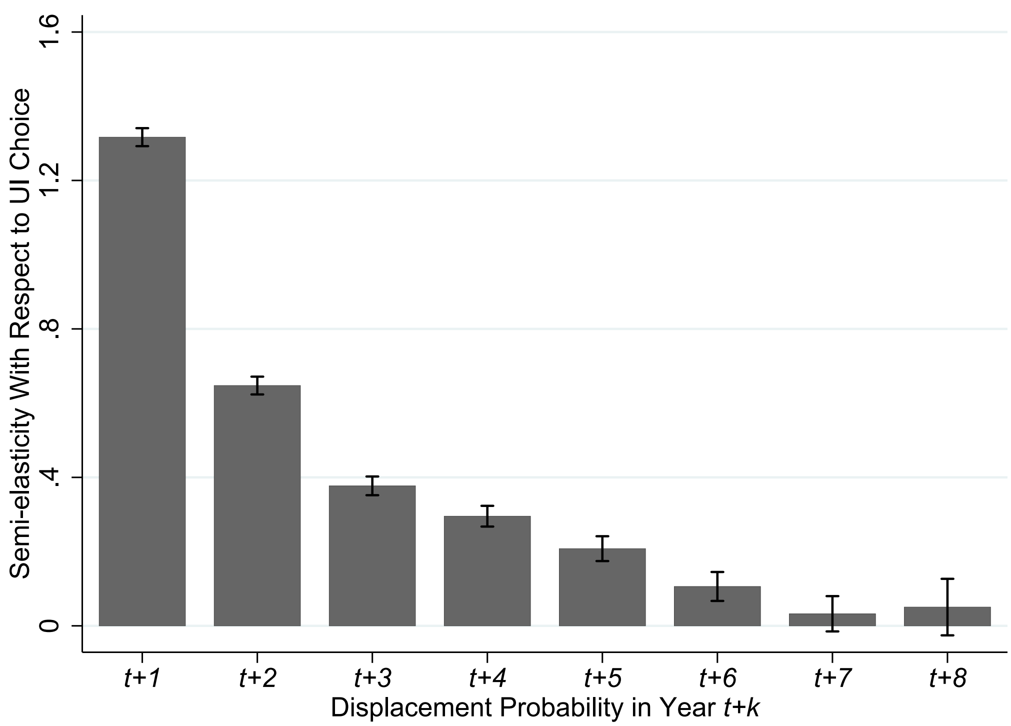
Risk Dynamics Our correlation tests use the risk outcomes in $t+1$, reflecting the idea that workers need to contribute for a year to be able to get the comprehensive coverage. However, the risk realization in $t+1$ may fail to fully capture the unemployment risk faced by an individual as she is making her coverage choice at time t, which justifies using risk realizations further into the future. In Figure B.2 we report the correlation of the insurance choice in t with displacement outcomes in $t+1$, $t+2$,... up to $t+8$. For each displacement outcome, the chart displays $\hat{\alpha_k}/\bar{\pi}$, that is the semi-elasticity of the realized risk outcomes in $t+k$ with respect to insurance choices in t, from a simple linear specification where we also control for all displacement outcomes in previous years ($t+k-1$, $t+k-2$, etc.):

$$\pi_{i,t+k} = \alpha_k D_i + Z_i' \beta + \sum_{l=0}^{k} \pi_{i,t+k-l} + \epsilon_i, \tag{21}$$

The first thing to note is that the estimated PCT for displacement risk in year $t+1$, using the linear specification (21) is equivalent to the PCT of Figure B.1 above, estimated from the non-linear specification (20). This is indicative that the PCT results are robust to functional form specifications.

The Figure also reveals an interesting dynamic pattern. The positive correlation between insurance choice and risk decreases rapidly as we consider displacement risk further in the future, but remains statistically significant up to six years. This pattern could indicate that workers' insurance choices incorporate private information about unemployment risk further into the future (albeit to a decreasing extent), but it may also be affected by moral hazard responses.

Figure B.2: Positive Correlation Tests - Dynamics



Notes: Risk realization in $t+1$ may fail to fully capture the unemployment risk faced by an individual as she is making her coverage choice at time t, which justifies using risk realizations for that individual further into the future. This Figure reports the correlation of insurance choice in t with displacement outcomes in $t+1$, $t+2$,... up to $t+8$. The Figure displays estimates of positive correlation tests following specification (12) estimated over the period 2002-2006. For each outcome, the chart displays $\hat{\alpha_k}/\bar{\pi}$, that is the semi-elasticity of the realized displacement rate in $t+k$ with respect to insurance choices in t. For each displacement outcome in year $t+k$, we control for displacement outcomes in previous years ($t+k-1$, $t+k-2$, etc.), for year fixed effects and for the limited set of characteristics Z that affect the unemployment insurance coverage available to individuals. See text for details.

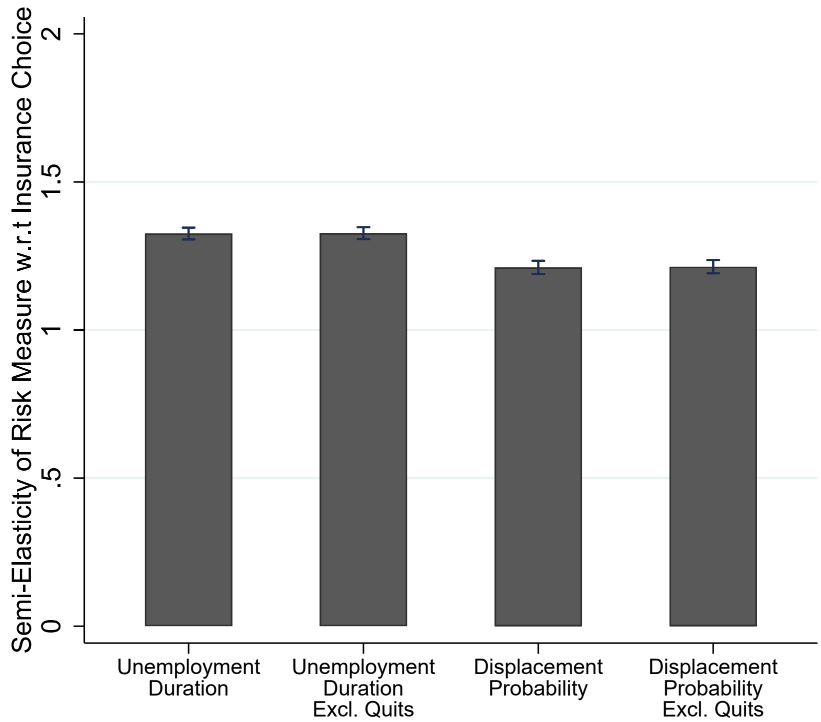
Unemployment Risk Excluding Involuntary Quits In the Swedish UI system, "quits following a valid reason" are eligible for unemployment benefits. They are therefore included in our measure of unemployment risk. The fact that involuntary quits are eligible to UI may raise the possibility of collusion between employers and employees to qualify actual voluntary quits as "quits following a valid reason". To understand to what extent this type of moral hazard drives the positive correlation between UI choices and realized unemployment risk, we exclude quits from the definition of unemployment. To do this, we use the fact that in the IAF data, a variable indicates whether an unemployment spell starts following a "quit for a valid reason".

We use again a simple linear specification:

$$\pi_i = \alpha D_i + Z_i'\beta + \epsilon_i, \tag{22}$$

In Figure B.3, we report $\hat{\alpha}/\bar{\pi}$, that is the semi-elasticity of the realized risk outcome in $t+1$ with respect to the insurance choice in t from this specification. We first use as an outcome π the total number of days spent unemployed in $t+1$, when including quits (first bar). Then, in the second bar, we report results where we use as an outcome π the total duration spent unemployed in $t+1$ when excluding involuntary quits from the definition of unemployment risk. We then replicate this exercize using as an outcome the probability of displacement in $t+1$ when including quits (third bar) and when excluding involuntary quits (fourth bar) from the definition of displacement risk. The figure shows that the positive correlation between unemployment risk and UI choices is almost unaffected by the inclusion or exclusion of involuntary quits.

Figure B.3: Positive Correlation Tests: Risk Outcomes Including and Excluding Involuntary Quits



Notes: This Figure reports the correlation of insurance choice in t with risk outcomes in $t+1$. The Figure displays estimates of positive correlation tests following specification (22) estimated over the period 2002-2006. For each outcome, the chart displays $\hat{\alpha}/\bar{\pi}$, that is the semi-elasticity of the realized risk with respect to insurance choices in t. For each outcome, we control for year fixed effects and for the limited set of characteristics Z that affect the unemployment insurance coverage available to individuals. See text for details.

B.2 Bivariate Probit & Non-parametric Tests

We now further investigate functional form restrictions and provide correct inference for the correlation tests.

First, we provide results of bivariate probit tests, popularized by Chiappori and Salanié [2000]. We specify both the choice of insurance coverage and the realization of our binary measure of unemployment risk (i.e., the probability of displacement) as probit models:

$$\begin{aligned} \mathbf{u}_i &= \mathbb{1}[Z'\alpha_1 + \epsilon > 0] \\ \pi_i &= \mathbb{1}[Z'\alpha_2 + \eta > 0] \end{aligned} \tag{23}$$

where $\mathbf{u}_i = u_{i,1} - u_{i,0}$ is the short-hand notation for the difference in indirect expected utility for individual i between being in the comprehensive plan and being in the basic plan. We allow for correlation ρ between the two error terms ϵ and η. The vector of controls Z contains the same variables as in specification (12). We provide in Table B.1 estimates of ρ and formal tests of the null that $\rho = 0$. Results confirm the presence of a strong and significant correlation between insurance choices and realized unemployment risk.

The functional forms involved in the bivariate probit tests are still restricted to the latent models being linear and the errors normal, excluding cross-effects or more complicated non-linear functions of the variables in Z. We therefore also produce results from non-parametric tests as in Chiappori and Salanié [2000]. The procedure of the test consists in partitioning the data into cells where all observations in a given cell have the same value for the variables in Z. The procedure then computes within each cell a Pearson's χ^2 test statistic for independence between $\mathbf{u}$ and π. This test statistic is asymptotically distributed as a $\chi^2(1)$ under the null hypothesis that $\mathbf{u}$ and π are statistically independent (within the cell). We report in the first column of Table B.2 results from this non-parametric procedure when cells are defined using the same controls Z as in specification (12) and where our risk measure π is the probability of displacement. Results again strongly confirm the presence of a positive correlation between insurance choices and unemployment probability. In Figure B.4 panel 1, we display the empirical distribution of the Pearson's χ^2 test statistics computed from all the cells to allow for comparison with a theoretical $\chi^2(1)$ distribution. Taking the largest absolute difference between the theoretical and the empirical distribution gives the Kolmogorov-Smirnov test statistic reported in Table B.2.

In columns (2) to (4) of Table B.2, we explore the robustness of the positive correlation test to adding more observable characteristics in the vector Z. In other words, we want to explore how much positive correlation would remain if the UI policy was allowed to differentiate coverage or prices along obvious observable dimensions that do not currently enter the UI policy schedule (such as age, gender, etc.). To this effect, we reproduce the non-parametric Kolmogorov-Smirnov test adding sequentially more observable characteristics to the vector Z when partitioning the data into cells. We start in column (2) of Table B.2 by adding demographic controls : age, then gender, and marital status. The Kolmogorov-Smirnov test statistic increases sharply, indicating that demographics may offer advantageous selection. Yet, we can still strongly reject the null

Table B.1: Positive Correlation Tests: Bivariate Probits

			Test $\rho = 0$	
	ρ	s.d.	χ^2	P-Value
Proba. of displacement	.3047	.0030	8842.4	0.00
Proba. of displacement excl. quits	.3056	.0031	8493.9	0.00

Notes: The Table reports positive correlation estimates between insurance and risk using bivariate probit models. We specify both the choice of insurance coverage and the probability of displacement as probit models allowing for correlation ρ between the two error terms ϵ and η. The Table reports estimates of ρ and its standard error. We also report results of formal tests of the null that $\rho = 0$. In the first row, we consider the probability of displacement. In the second row we consider the probability of displacement excluding quits, as some quits may be eligible for UI after a waiting period. See text for details.

of no positive correlation between risk and insurance choice. In column (3), we add controls for education (four categories), and industry (1-digit code). The Kolmogorov-Smirnov test statistic does not seem to be affected much by the inclusion of these controls for skills and other labor market characteristics. In column (4), we finally add controls for past unemployment history (dummies for having been unemployed in $t-1$, $t-2$ and up to $t-8$). The Kolmogorov-Smirnov test statistic decreases as a result, suggesting that past unemployment history creates significant adverse selection.

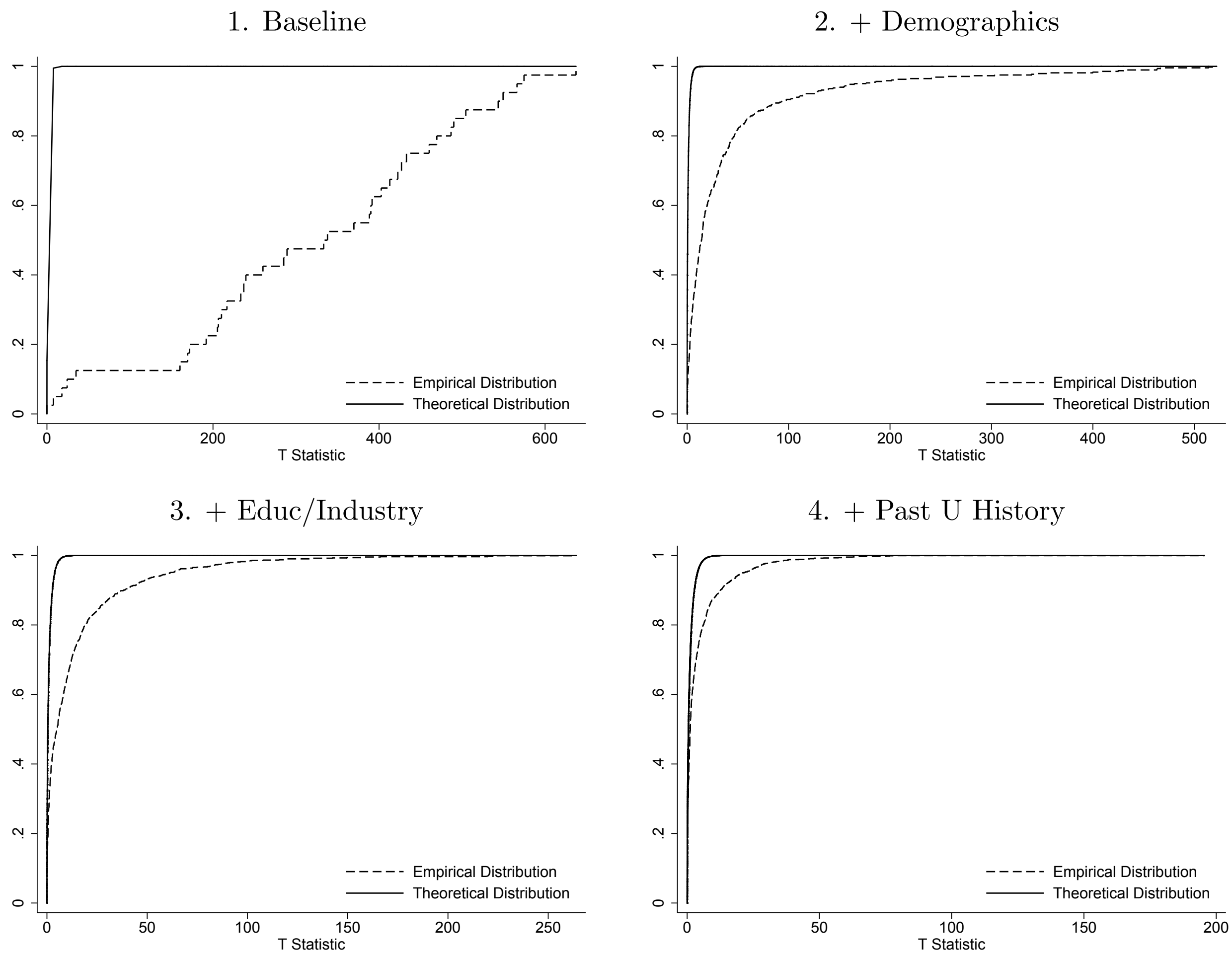
For all specifications of columns (1) to (4) of Table B.2, the corresponding panels 1 to 4 of Figure B.4 display the empirical distribution of the Pearson's χ^2 test statistics computed from all the cells to allow for comparison with a theoretical $\chi^2(1)$ distribution.

Table B.2: Positive Correlation Tests: Non-Parametric Tests

	(1)	(2)	(3)	(4)
	Variables included in partitioning the data in cells			
	Baseline	+ Demographics	+ Educ & Industry	+ Past U History
# of cells	40	484	1,124	1,923
Average cell size	50,903	3,181	958	415
Median cell size	35,275	1,270	346	141
Minimum cell size	14,202	88	6	5
Fraction of cells too granular	0%	24%	65%	80%
Fraction of rejected cells	98%	74%	53%	28%
Kolmogorov-Smirnov stat.	5.98	15.37	16.20	10.47
Binomial p-value	0%	0%	0%	0%

Notes: The Table reports results from non-parametric tests of correlation between insurance choices in t and probability of displacement in $t+1$. The procedure of the test consists in partitioning the data into cells where all observations in a given cell have the same value for the variables in Z. Columns (1) to (4) differ in the control variables included in Z and used to partition the data. The procedure then computes within each cell a Pearson's χ^2 test statistic for independence between $\mathbf{u}$ and π. This test statistic is asymptotically distributed as a $\chi^2(1)$ under the null hypothesis that $\mathbf{u}$ and π are statistically independent (within the cell). The critical values of this statistic for 95% and 99% confidence are 1.36 and 1.63 respectively. The reported Kolmogorov-Smirnov test statistic is scaled by $\sqrt{n}$ where n is the number of cells. When adding a lot of controls to the vector Z, some cells can become too granular to compute the test statistic (division by zero). We therefore also report in the Table the number of cells that are too granular.

Figure B.4: Positive Correlation Tests - Distribution of χ^2 test statistics from all cells vs Theoretical $\chi^2(1)$ distribution - Additional Controls



Notes: The Figure displays the empirical distribution of the Pearson's χ^2 test statistics for independence between $\mathbf{u}$ (UI choices) and π, the probability of layoff in $t+1$, computed from all the cells where we split individuals in cells corresponding to various observable characteristics. In panel 1, we only use priced characteristics (baseline controls of the positive correlation tests), corresponding to the test implemented in column (1) of Table B.2. In panel 2, we add controls for demographics (cf. column (2) of Table B.2). Panel 3 and 4 add education, industry and past unemployment history controls (cf. column (3) and (4) of Table B.2). The χ^2 test statistic is asymptotically distributed as a $\chi^2(1)$ under the null hypothesis that $\mathbf{u}$ and π are statistically independent (within the cell). We therefore compare this distribution with a theoretical $\chi^2(1)$ distribution. Taking the largest absolute difference between the theoretical and the empirical distribution gives the Kolmogorov-Smirnov test statistic reported in Table B.2.

Table B.3: Summary Statistics - By Coverage

	A. Under Basic				**B. Under Comprehensive**			
	Mean	P10	P50	P90	Mean	P10	P50	P90
	I. Unemployment							
Displacement probability	1.96%	-	-	-	3.21%	-	-	-
Displacement probability (exc. quits)	1.81%	-	-	-	3%	-	-	-
Unemployment probability	2.29%	-	-	-	3.85%	-	-	-
Days unemployed	2.84	0	0	0	5.65	0	0	0
Predicted days unemployed under Basic	2.96	1.89	2.58	3.68	3.66	2.18	2.74	4.09
Predicted days unemployed under Comprehensive	5.34	3.45	4.67	6.82	5.91	3.66	4.78	7.24
Unemployment spell duration (days)	137.57	26	90	283	148.26	22	91	307
Fraction receiving layoff notification	.04	-	-	-	.06	-	-	-
Fraction switching firms	.1	-	-	-	.09	-	-	-
	II. Union and UI Fund Membership							
Union membership	.13	-	-	-	.84	-	-	-
Switch from coverage 0 to 1	-	-	-	-	.02	-	-	-
Switch from coverage 1 to 0	.01	-	-	-	-	-	-	-
	III. Demographics							
Age	35.52	25	33	55	41.7	27	42	55
Years of education	12.97	11	12	16	12.84	11	12	16
Fraction men	.63	-	-	-	.51	-	-	-
Fraction married	.32	-	-	-	.46	-	-	-
	IV. Income and Wealth, SEK 2003(K)							
Gross earnings	233.8	65.3	186.7	416	251.5	115	234.9	385.4
Net wealth	649.6	-195.1	25.7	1521.7	343.5	-155.5	102.4	1083.5
Bank holdings	73.9	0	0	135.8	45.3	0	0	120.4
N	2,296,727				15,003,779			

Notes: The Table breaks down the summary statistics by UI coverage for our main sample of interest over the period 2002 to 2006. See Table 1.

Appendix C Impact of Predictable Risk on UI Choice: Further Evidence

In this appendix we present further evidence regarding the relationship between predictable risk and UI choices:

(i) we present non-parametric evidence of the relationship linking UI choices to both $\hat{\pi}_0$ and $\hat{\pi}_1$

(ii) we present quasi-experimental evidence showing how the various institutional risk shifters detailed in Appendix A, which enter our predicted risk model, separately affect selection into coverage.

C.1 Non-Parametric Evidence on the Relationship between Predicted Risk and Insurance Choice

The positive correlation tests between predicted risk and UI choice from section 4.2 shows clearly that individuals buying the supplemental coverage have a higher predictable risk on average in both the basic coverage ($E_1[\hat{\pi}_0] > E_0[\hat{\pi}_0]$) and the comprehensive coverage ($E_1[\hat{\pi}_1] > E_0[\hat{\pi}_1]$). We provide here more detailed non-parametric evidence on the relationship between insurance choice and predicted risk in both coverages.

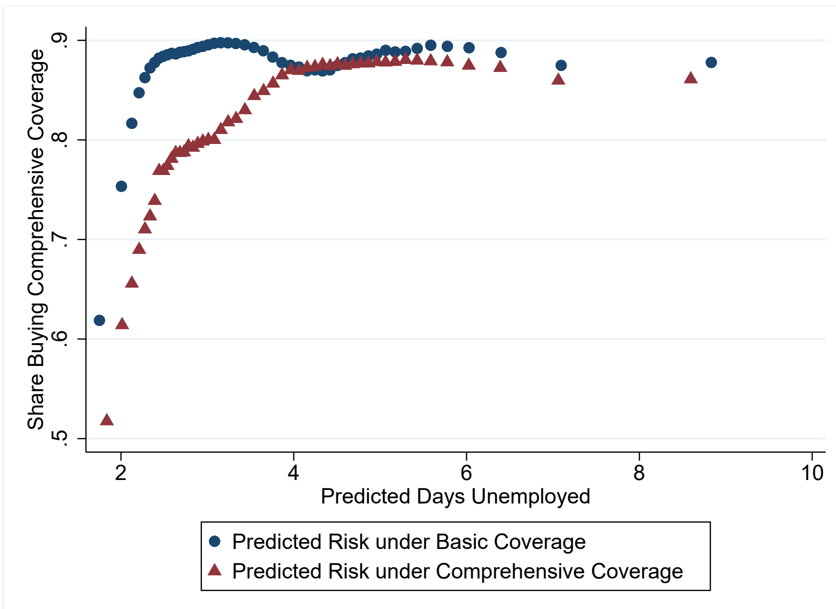
For this purpose, Figure C.1 offers a bin scatter correlating the probability to buy the comprehensive UI coverage in year t with the predicted number of days unemployed of individual i, respectively under the basic coverage $\hat{\pi}_0$ and under the comprehensive coverage $\hat{\pi}_1$, based on her observable characteristics year t. The graph confirms evidence from the positive correlation tests of a strong positive correlation between individuals' predictable risk and their probability to buy the comprehensive UI coverage. Interestingly, the graph also suggests that the strong positive correlation between risk and insurance coverage is mostly driven by what happens at the bottom of the predicted risk distribution. Only about a half of individuals at the bottom of the predicted risk distribution ($\hat{\pi}_1 < 2$ days) buy the comprehensive coverage. But this fraction quickly rises as the predicted risk increases. It is then very stable, at around 85 to 90% for individuals with predicted risk $\hat{\pi}_1 < 5$ days. Note finally that conditional on the fraction buying the comprehensive coverage, the difference between predicted risk under basic and under comprehensive coverage captures the presence of moral hazard.

C.2 Risk Shifter & UI Choices I: Average Firm Layoff Risk

The previous evidence focuses on risk measures from our predicted risk model. This model folds all sources of variations of observable risk together into a unique measure of predictable risk. We now also shed light on how the various institutional risk shifters that enter the predicted risk model individually affect selection into coverage.

The first source of risk variation is average firm level risk. We define again the average firm displacement risk $\bar{\pi}_{-i,j}$ of worker i working in firm j as the average probability of displacement

Figure C.1: Predicted Risk and UI Coverage Choice



Notes: The Figure displays a bin scatter correlating the probability to buy the comprehensive UI coverage in year t with the predicted number of days unemployed of individual i, respectively under the basic coverage $\hat{\pi}_0$ and under the comprehensive coverage $\hat{\pi}_1$, based on her observable characteristics year t. The measures of predictable unemployment risk under basic and comprehensive coverage are from the model presented in Section 2.3. The model combines flexibly all observable sources of risk, including institutional shifters of risk such as the full history of the firm layoff notifications, and the relative tenure ranking of the individual. Model selection is based on the Lasso approach for zero-inflated poisson suggested by Banerjee et al. [2018]. To allow for moral hazard, we estimate a model of risk for individuals under the basic coverage, and a separate model of risk for individuals under the comprehensive coverage. The model predicts the number of days spent unemployed in year $t+1$ based on observable characteristics in year t.

of all other workers within firm j excluding individual i over all years where the firm is observed active in our sample years.

In section Appendix A we showed that there is significant heterogeneity in these average firms' separation rates, and that individuals' unemployment risk is very strongly correlated with this average firm level risk (panel A of Figure A.1).

We now investigate how average firm level risk correlates with unemployment insurance choices.

Cross-Sectional Evidence The first strategy consists in simply using the cross-sectional variation in displacement risk across firms. In Figure C.2 panel A, we group individuals in 50 equal size bins of firm layoff risk, and plot their average firm layoff risk against their average probability of buying supplemental coverage, residualized on the same vector Z of baseline controls affecting UI contracts used in the positive correlation test of Section 4.1.

The graph displays a strong positive correlation between firm layoff risk and individuals' probability to buy the comprehensive UI coverage

We then estimate the correlation between average firm level risk $\bar{\pi}_{-i,j}$ and willingness-to-pay by running the following two-stage least square specification:

$$\begin{aligned} D_i &= \beta_{2SLS} \cdot \pi_i + Z_i'\alpha_1 + \epsilon \\ \pi_i &= \zeta \cdot \bar{\pi}_{-i,j} + Z_i'\alpha_2 + \eta \end{aligned} \tag{24}$$

where D_i is our indicator variable for buying the supplemental coverage. This specification instruments individual realized risk by the average firm layoff risk and therefore exploits only variation in predictable risk coming from average firm layoff risk. For useful comparison, we also report the coefficient estimate β_{OLS} of the following OLS specification correlating D with individual risk:

$$D_i = \beta_{OLS} \cdot \pi_i + Z_i'\alpha + \nu \tag{25}$$

We estimate both models on our baseline sample of workers pooling all observations for years 2002 to 2006. We use as a measure of realized risk π_i the realized displacement risk excluding quits in year $t+1$. We find a positive and strongly significant coefficient $\beta_{2SLS} = .50$ (.01) indicating that workers who work in firms that exhibit higher turnover rates are significantly more likely to buy the comprehensive coverage.

We also find that β_{2SLS} is much larger than β_{OLS}, which is also informative. Clearly, the two-stage least square procedure removes potential attenuation bias from measurement error in β_{OLS}. But the two-stage least square, by projecting choices only on the average firm layoff dimension of displacement risk introduces some potential selection, if $\text{Cov}(\bar{\pi}_{-i,j}, \epsilon) \neq 0$. In other words, if workers who self-select into riskier firms are different along observed or unobserved characteristics correlated with willingness-to-pay for insurance, β_{2SLS} will capture this additional selection effect.

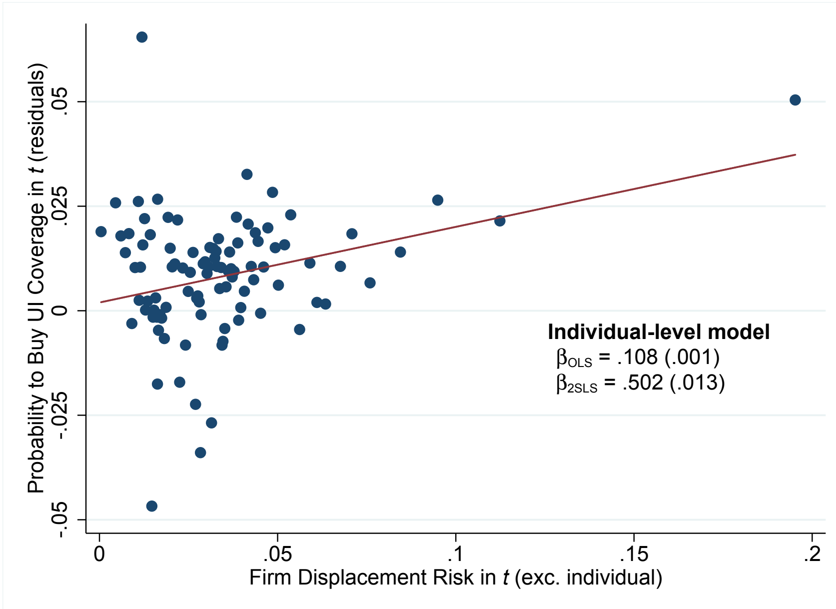
In panel B of Figure C.2, we explore the importance of such selection along observable characteristics in explaining the magnitude of β_{2SLS}. We introduce in the vector Z of specifications (24) and (25) a rich set of additional controls: age, gender, marital status, education (four categories), industry (1-digit code), occupation (1-digit code), wealth level (quartiles) and past unemployment history (dummies for having been unemployed in $t-1$, $t-2$ and up to $t-8$). We still find a strong positive correlation between insurance choices and firm layoff risk ($\beta_{2SLS} = .245$). But adding these controls decreases the magnitude of the correlation between risk and UI choices significantly.

Even with this rich set of controls, β_{2SLS} might still be picking some correlation between average layoff risk and *unobserved* characteristics affecting UI choices. This will be the case if workers who select to work in riskier firms have different preferences for insurance and/or if the there is an unobserved effect of riskier firm environments on insurance choices: firms with high turnover may have different prevalence of collective bargaining, different firm cultures that can affect individuals' UI choices for instance.

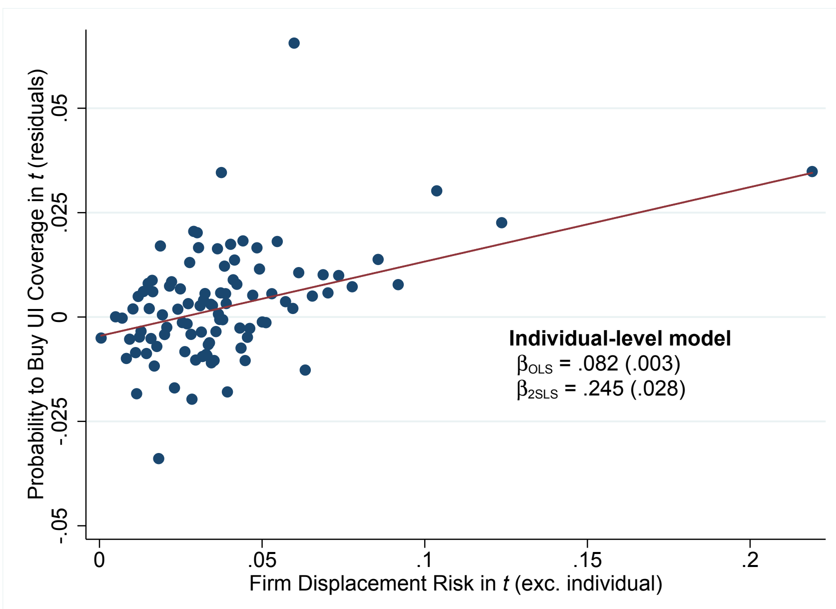
Decomposing the error term in specification (24) $\epsilon = \kappa_i + \rho_j$ into an individual specific component κ_i and a firm specific component ρ_j, we can think of the the selection introduced by average layoff risk as the combination of individual fixed effects and firm fixed effects. We first move to a firm switcher design that allow us to control more directly for the unobserved individual specific component κ_i. In subsection C.3 we then show how to deal with both the individual specific (κ_i) and firm specific (ρ_j) sources of potential selection.

Figure C.2: Firm Level Risk and UI Coverage Choice

A. Baseline Controls for Contract Space



B. With Additional Demographic Controls



Notes: The Figure uses cross-sectional variation in displacement risk across firms as a risk shifter to estimate how UI coverage choices react to variation in risk that is not driven by individual moral hazard. Panel A groups individuals in 50 equal size bins of firm layoff risk, and plot their average firm layoff risk against their average probability of buying supplemental coverage, residualized on the same vector X of baseline controls affecting UI contracts used in the positive correlation test of Section 4.1. We report on the graph the coefficient β_{OLS} from an OLS regression of specification (25) and then the estimated coefficient β_{2SLS} from our two-stage least square model (24) where we use $Z = \pi_{-i,j}$ as a risk shifter. In panel B, we replicate the same procedure, but now add to the regression the same rich set of demographic controls used in Figure 6, and find a similar strong positive correlation between insurance choices and firm layoff risk.

Firm Switcher Design In this strategy, we use the panel dimension of the data to control for the selection introduced by individual specific heterogeneity κ_i.

To this end, we focus on individuals who change firms ("switchers"). The employer-employee matched data ($RAMS$) registers all existing labor contracts on a monthly basis. We define a switch as moving from having a labor contract with firm j (the origin firm) to having a contract with firm k (the destination firm), without any recorded non-employment spell between these two contracts. We focus on individuals with more than 1 year of tenure in the origin firm. Switchers experience a change in their layoff risk coming from underlying variation in two risk shifters: their tenure ranking changes, and so does their underlying firm layoff risk.

First, switchers experience a reduction in their relative tenure ranking, as they become the "last-in" when they move to the destination firm. To document the magnitude of the variation in relative tenure ranking and corresponding layoff risk, following a firm switch, we define year $n = 0$ as the year of a firm switch, and run, on the sample of firm switchers, event studies of the form:

$$T_{i,n} = \sum_k \delta_k \cdot \mathbb{1}[n = k] + \mathbf{Z_i'}\alpha + \epsilon_{i,n} \tag{26}$$

where T_n denotes the tenure ranking of individual i in event year n, $\mathbb{1}[n = k]$ are a set of event time dummies, and Z is the vector of baseline controls affecting UI contracts defined in section 4.1. Figure C.3 displays the evolution of relative tenure ranking of switchers as a function of event time by plotting the coefficients δ_k, taking event time $n = -1$ as the omitted category. The graph confirms that relative tenure ranking decreases sharply at the moment of the firm switch.

Figure C.4 panel A explores how this variation in relative tenure ranking affects the probability of displacement over event time n. To this effect, we estimate a similar event study specification as in (26) where we use the probability of displacement π_i in year $t + 1$ as an outcome. The graph shows that the displacement rate one year ahead increases sharply and significantly at the time of the firm switch.

In Figure C.5 panel A, we run a similar event study specification with D_i, a dummy for buying the comprehensive UI coverage as an outcome. The figure shows that the probability of buying the comprehensive coverage increases sharply by about 2.2 percentage points at the time of the firm switch. On the graph, we also display the coefficient from the following two-stage least square specification:

$$\begin{aligned} D_{i,n} &= \kappa_i + \beta_{2SLS} \cdot \pi_{i,n} + \mathbf{Z'_{i,n}}\alpha_{\mathbf{1}} + \epsilon_{i,n} \\ \pi_{i,n} &= \nu_i + \zeta \cdot \mathbb{1}[n \geq 0] + \mathbf{Z'_{i,n}}\alpha_{\mathbf{2}} + \eta_{i,n} \end{aligned} \tag{27}$$

where we use a dummy for having switched firm ($\mathbb{1}[n \geq 0]$) as risk shifter for individual displacement probability $\pi_{i,n}$ and control for individual fixed-effects. This specification is estimated on the sample of all workers who ever experience a firm switch between 2002 and 2006 and who have more than 1 year of tenure in the origin firm. β_{2SLS} is positive and strongly significant, which again indicates that the positive correlation tests are not simply picking up moral hazard responses to insurance coverage.

While these event study specifications control for fixed underlying heterogeneity across individuals that may affect their UI choices (κ_i), one concern with this original implementation of the firm switchers design is that individuals are somewhat inert, and decide to reoptimize their UI choices only at specific times, like, for instance, when they switch firm.

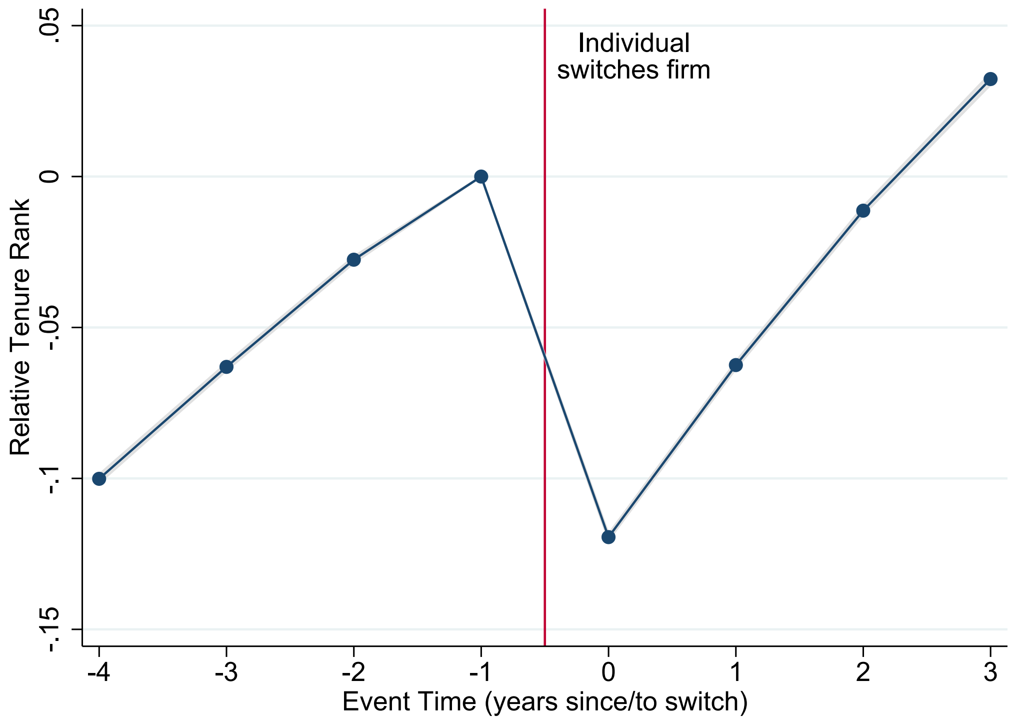
To mitigate the concern that the surge in UI coverage at the time of the switch is the result of the specific timing of UI choices and not a response to the change in underlying risk, we exploit additional variation in risk in the switchers design coming from changes in underlying firm layoff risk. While all switchers experience an increase in their displacement probability due to the decline in their tenure ranking, the effect of a switch on individual displacement probability exhibits large differences according to whether their destination firm is much riskier ("positive shock") or a lot less risky ("negative shock") than their origin firm. We therefore split the population of switchers according to their rank in the distribution of $\Delta_{j,j'}\bar{\pi}_{-i} = \bar{\pi}_{-i,j'} - \bar{\pi}_{-i,j}$, the change in their underlying average firm layoff risk when moving from firm j to firm j'. In Figure C.4 panel B, we contrast individuals in the bottom decile of $\Delta_{j,j'}\bar{\pi}_{-i}$ (large negative shock, i.e., individuals who experience a large negative decline in their firm layoff risk, going from a high risk to a low risk firm), and individuals in the top decile of $\Delta_{j,j'}\bar{\pi}_{-i}$ (large positive shock, i.e., individuals who experience a large increase in their firm layoff risk going from a low risk to a high risk firm). The Figure confirms that individuals experiencing a large positive shock in their firm layoff risk exhibit a significantly larger increase, of about 2 percentage point, in their displacement probability at the time of the switch, relative to individuals experiencing a large negative shock.

In panel B of Figure C.5, we now compare the evolution of insurance choices around firm switch for individuals experiencing large positive vs large negative shocks. We run event study specification (26) with D_i, a dummy for buying the comprehensive UI coverage as an outcome, separately for the sample of individuals experiencing large positive shocks and for the sample of individuals experiencing large negative shocks. The graph indicates that the increase in the probability to buy UI around firm switch is significantly larger (by about 1.5 percentage point) among individuals moving to significantly more risky firms relative to those moving to less risky firms. We also report on the graph the estimated coefficient $\beta_{2SLS} = .57$ (.08) of the two-stage model:

$$\begin{aligned} D_{i,n} &= \kappa_i + \beta_{2SLS} \cdot \pi_{i,n} + \textstyle\sum_k \delta_k \cdot \mathbb{1}[n=k] + \mathbf{Z'_{i,n}}\alpha_{\mathbf{1}} + \epsilon_{i,n} \\ \pi_{i,n} &= \nu_i + \zeta \cdot \mathbb{1}[n \geq 0] \cdot \Delta\bar{\pi}_{-i,j} + \mathbf{Z'_{i,n}}\alpha_{\mathbf{2}} + \eta_{i,n} \end{aligned} \tag{28}$$

This model uses firm switch interacted with the change in average firm level layoff risk $\Delta\bar{\pi}_{-i,j}$ as risk shifter for individual displacement probability. This model estimated on the sample of all workers who ever experience a firm switch between 2002 and 2006 and who have more than 1 year of tenure in the origin firm. The results suggest that the probability to buy the comprehensive coverage is strongly correlated with average firm layoff risk, even after controlling for individual unobserved heterogeneity with this switcher design strategy.

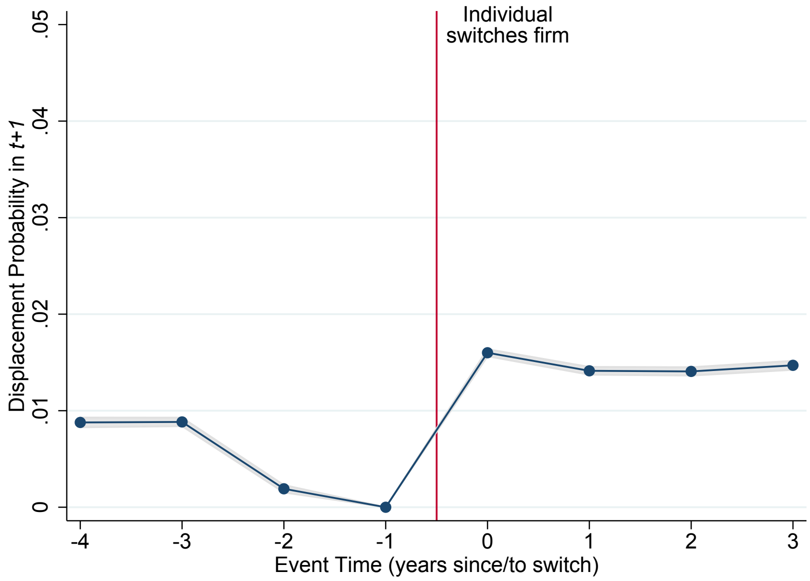
Figure C.3: Switchers Design: Relative Tenure Ranking as a Function of Event Time



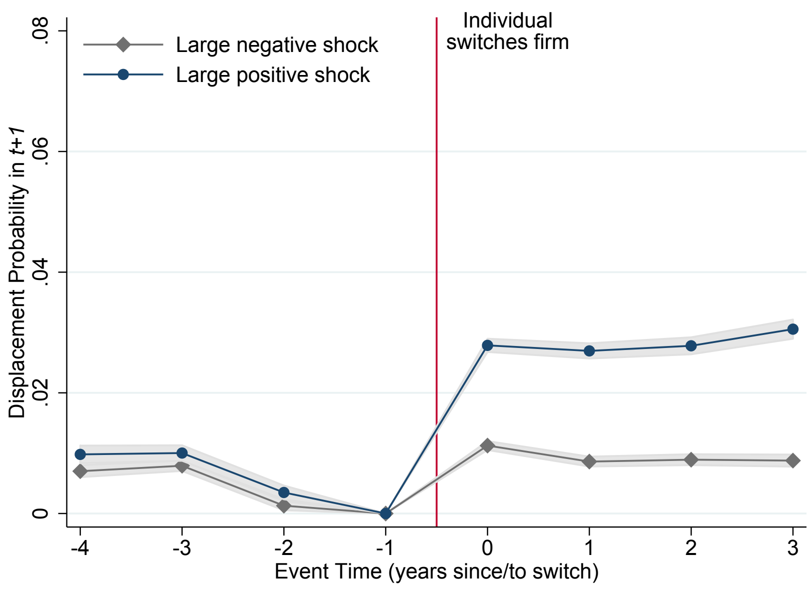
Notes: The Figure focuses on “firm switchers”, i.e. individuals moving from having a labor contract with firm j to having a contract with firm k, without any recorded non-employment spell between these two contracts. We focus on individuals with more than 1 year of tenure in the origin firm. In this Figure we show that switchers experience a variation in their layoff risk coming from underlying variation in their relative tenure ranking. Relative tenure ranking affects displacement probability due to the strict enforcement of the Last-In-First-Out (LIFO) principle in Swedish labor laws. To follow the rules pertaining to the application of LIFO, relative tenure ranking is defined within each establishment times occupation group using the RAMS employer-employee data since 1985. The chart displays estimates of the event study specification (26) using relative tenure ranking as an outcome. The graph shows that relative tenure ranking drops abruptly at the time of the firm switch. Panel A of Figure C.4 shows that this drop in tenure ranking translates in a significant increase in displacement risk.

Figure C.4: Firm Switchers - Displacement Rate in $t+1$ as a Function of Time to/since Firm Switch

A. All Switchers



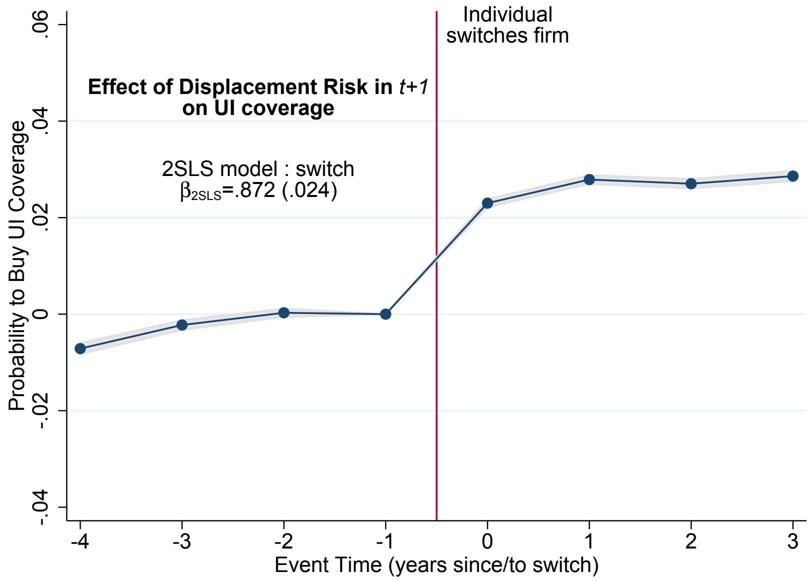
B. Switchers Experiencing Large Positive Firm Layoff Risk Shock vs Large Negative Firm Layoff Risk Shock



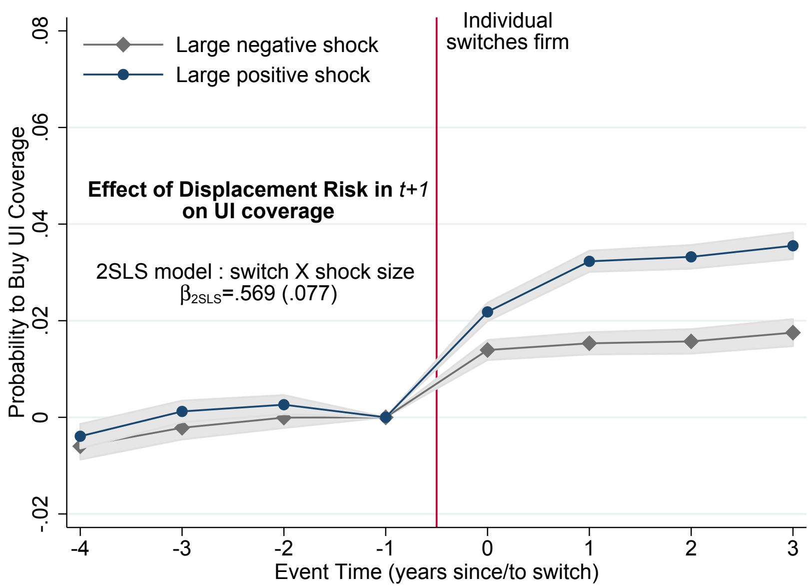
Notes: The Figure focuses on "firm switchers", i.e. individuals moving from having a labor contract with firm j to having a contract with firm k, without any recorded non-employment spell between these two contracts. We focus on individuals with more than 1 year of tenure in the origin firm. Switchers experience a variation in their layoff risk coming from underlying variation in both risk shifters: their tenure ranking changes, and so does their underlying firm layoff risk. In panel A, we display estimates of the event study specification (26) using displacement risk in $t+1$ as an outcome. The graph shows that the displacement risk increases sharply and significantly at the time of the firm switch. In panel B, we split the population of switchers according to their rank in the distribution of $\Delta_{j,j'}\pi_{-i} = \pi_{-i,j'} - \pi_{-i,j}$, the change in their underlying firm risk when moving from firm j to firm j'. We focus on individuals in the bottom decile of $\Delta_{j,j'}\pi_{-i}$ (large negative shock, i.e., individuals going from a high risk to a low risk firm), and individuals in the top decile of $\Delta_{j,j'}\pi_{-i}$ (large positive shock).

Figure C.5: Firm Switchers - UI Coverage Choices as a Function of Time to/since Firm Switch

A. All Switchers



B. Switchers Experiencing Large Positive Firm Layoff Risk Shock vs Large Negative Firm Layoff Risk Shock



Notes: The Figure focuses on "firm switchers". In panel A, we display estimates of the event study specification (26) using UI coverage V as an outcome. The Figure shows that the probability of buying the comprehensive coverage increases sharply at the time of the firm switch. In panel B, we split the population of switchers according to their rank in the distribution of $\Delta_{j,j'}\pi_{-i} = \pi_{-i,j'} - \pi_{-i,j}$, the change in their underlying firm risk when moving from firm j to firm j', as in Figure C.4 panel B. The graph indicates that the increase in the probability to buy UI around firm switch is significantly larger among individuals moving to significantly more risky firms relative to those moving to less risky firms. On both panels, we display the coefficient from a two-stage least square fixed-effect specification similar to (24) where we use firm switch (and firm switch interacted with shock size) as risk shifter Z for individual displacement probability.

C.3 Risk Shifter & UI Choices II: Layoff Notifications and LIFO

The previous section suggests a strong correlation between firm layoff risk and UI choices, indicative of the presence of significant adverse selection. As explained above though, firm layoff risk may be correlated with willingness-to-pay for UI, either through unobserved individual specific heterogeneity (κ_i) or unobserved firm specific heterogeneity (ρ_j). The firm switcher design above deals with individual specific heterogeneity (κ_i), but may still pick up selection on firm level heterogeneity ρ_j if firm heterogeneity is correlated with $\Delta\bar{\pi}_{-i,j}$.

We now show how layoff notifications and the application of the Last-In-First-Out (LIFO) principle enables to identify the effect of predictable risk on UI choices controlling jointly for firm level heterogeneity ρ_j and individual level heterogeneity κ_i. We leverage the fact that layoff notifications and LIFO creates variation in layoff risk both within firm and across individuals over time.

In section 2.1, we described the institutional details of the layoff notification system and its interaction with the LIFO rule. We also explained and demonstrated in Appendix A, that layoff notifications signal a significant change in a firm layoff risk. In particular, we reported in Figure A.2 that the displacement probability of workers increases sharply and significantly around the first layoff notification event in the history of the firm. We also showed in Figure A.1 panel B that the effect of a layoff notification on displacement probability is strongly heterogenous depending on the relative tenure ranking of workers. Workers with relative tenure ranking below .5 have a much higher probability of being laid-off following a layoff notification than workers with relative tenure ranking above .5.

We now show how UI choices correlate with this variation in risk stemming from the interaction between a notification event and relative tenure ranking. We follow the same event study empirical approach as in section Appendix A around the event of a layoff notification. We define event year n as the year to/since the firm emits its first layoff notification.

Our sample is the panel of workers who are employed in the firm at the date this layoff notification is emitted to the PES. All these workers constitute our treatment group. We follow, as in Appendix A a matching strategy and create a control panel of workers. To do this, we use nearest-neighbor matching to select a set of firms that are similar, along a set of observable characteristics, to the firms emitting a layoff notification, but never emit a layoff notification.[49] We allocate to the matched firm in the control group a placebo event date equal to the layoff notification date of her nearest-neighbor in the treated group of firms. We then select workers that are in the control firm at the time of the placebo event date to create our matched panel of control individuals.

In Figure C.6 we split the sample by tenure ranking at the time the layoff notification is emitted and report the evolution of the average fraction of individuals buying the supplemental coverage in our treatment group and in the matched control group.[50] Panel A of Figure C.6 reports the

[49] The covariates used for matching are the number of employees, the 4 digit sector codes of the firm, the average earnings and average years of education of workers in the firm.

[50] For control workers we use their tenure ranking at the time of the placebo layoff notification.

evolution of the fraction buying the supplemental UI coverage for workers with relative tenure ranking below 50% in year $n = 0$. The graph shows that UI coverage increases significantly among the treated group, starting one year before the layoff notification is sent, which suggests the existence of some degree of private information among workers regarding the timing of the layoff notification. In panel B, we report the evolution of UI choices for the sample of workers with relative tenure ranking above 50% in year $n = 0$. The graph displays no sign of variation in the fraction of individuals buying the comprehensive coverage around the notification event.

On both panels, we also report estimates $\hat{\beta}$ of the reduced form specification:

$$D_{i,n} = \kappa_i + \rho_j + \beta \cdot \mathbb{1}[n \geq 0] \cdot \mathbb{1}[T = 1] + \theta \cdot \mathbb{1}[n \geq 0] + \mathbf{Z'_{i,n}}\alpha_{\mathbf{1}} + \epsilon_{i,n} \tag{29}$$

as well as estimates $\hat{\beta}_{2SLS}$ from the following two-stage specification:

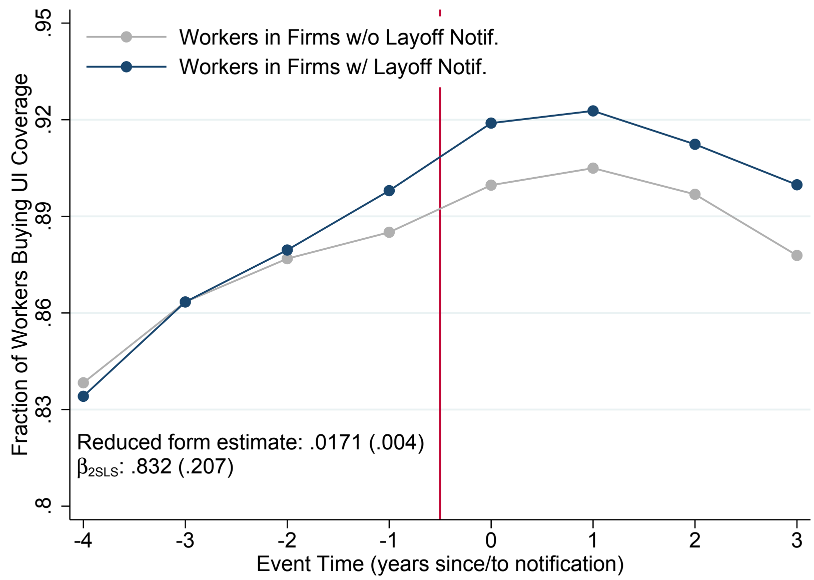
$$\begin{aligned} D_{i,n} &= \kappa_i + \rho_j + \beta_{2SLS} \cdot \pi_{i,n} + \textstyle\sum_k \delta_k \cdot \mathbb{1}[n = k] + \mathbf{Z'_{i,n}}\alpha_{\mathbf{1}} + \epsilon_{i,n} \\ \pi_{i,n} &= \nu_i + \gamma_j + \zeta \cdot \mathbb{1}[n \geq 0] \cdot \mathbb{1}[T = 1] + \theta \cdot \mathbb{1}[n \geq 0] + \mathbf{Z'_{i,n}}\alpha_{\mathbf{2}} + \eta_{i,n} \end{aligned} \tag{30}$$

The above two-stage model specification uses variation in risk stemming from being in a firm having emitted a layoff notification, and controls for both individual fixed effects (κ_i) and firm fixed effects (ρ_j). The comparison between the estimates for the low vs high tenure ranking sample further exploits the additional layer of variation in displacement risk coming from the interaction between a notification event and relative tenure ranking. Results show that individuals with low tenure ranking strongly respond to the variation in risk arising from a layoff notification and are significantly more likely to buy the comprehensive coverage as a result: $\beta_{2SLS} = .84$ (.21). To the contrary, the UI choices of individuals with high tenure ranking do not significantly respond to a layoff notification.

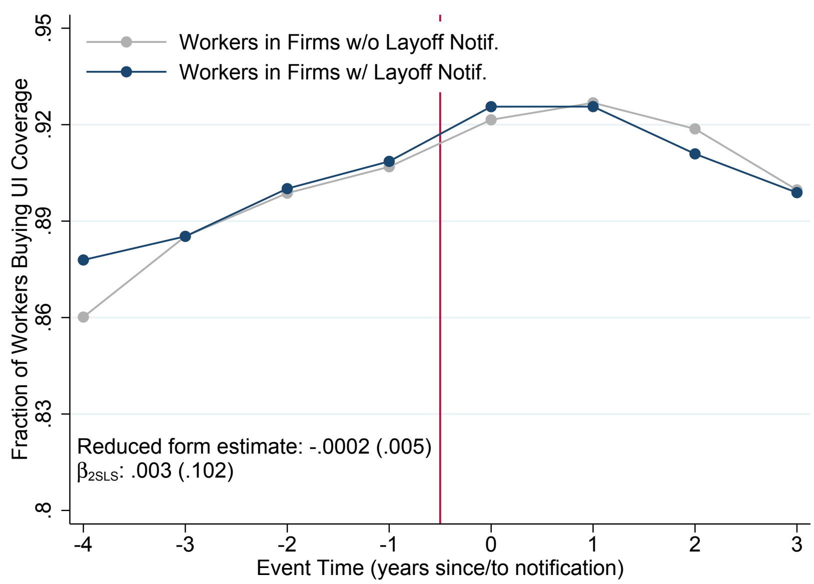
Summary of evidence Taken together, the evidence from this appendix strongly suggests that UI choices do significantly respond to the various sources of variations in individuals' predictable unemployment risk. The different strategies clearly differ in terms of the way they control for underlying selection on unobserved heterogeneity into the comprehensive coverage, as well as in terms of the population of compliers. Yet, we systematically find a strong positive and significant relationship between the probability to buy the comprehensive coverage and the observable risk shifters entering our predicted risk model. This overall confirms that the strong correlation between predictable risk and UI choices documented in section 4.2 does capture the presence of significant adverse selection into the comprehensive coverage.

Figure C.6: Layoff Notification

A. Workers with Relative Tenure Ranking < .5 at Event Time 0



B. Workers with Relative Tenure Ranking ≥ .5 at Event Time 0



Notes: The Figure uses layoff notification events interacted with relative tenure ranking as a source of variation in displacement risk to investigate how UI coverage choices react to variations in underlying risk. Panel A reports the evolution of UI coverage around the time of the first layoff notification for the panel of workers in the treated group and for workers in our placebo (control) group, restricting the sample to workers with relative tenure ranking below 50% in year $n = 0$. The Figure shows that UI coverage increases significantly among the treated group, starting one year before the layoff notification is sent, which suggests the existence of some degree of private information among workers regarding the timing of the layoff notification. In panel B, we report similar estimates but for the sample of workers with relative tenure ranking above 50% in year $n = 0$. The graph displays no sign of variations in individuals insurance coverage among the event. On both panels, we display the estimated coefficient β_{2SLS} of our two-stage least square model using the layoff event interacted with tenure and a dummy for being in the treatment group as a risk shifter Z.

Appendix D Price Variation: Additional Material

In this appendix we present additional results using the 2007 price variation to identify adverse selection:

(i) we present further non-parametric evidence of adverse selection using additional risk outcomes

(ii) we show how adverse selection would survive the inclusion of many unused demographic observables in the Swedish UI policy

(iii) we provide evidence showing that our ranking of willingness-to-pay for the comprehensive coverage correlates strongly with proxies for the value of unemployment insurance and for risk preferences.

(iv) we address potential concerns, such as inertia, to the validity of our ranking of individuals by willingness-to-pay.

Alternative risk outcomes In our baseline analysis of the 2007 reform in section 5, we use total number of days unemployed in 2008 as our main outcome. Here, we show that our estimates of adverse selection are again robust to using alternative risk outcomes. We look at the displacement probability in t+1, t+2,... up to t+5. To control for observables Z, we model the probability of displacement as a probit:

$$E(\pi|Z) = \Phi(Z'\beta + \sum_j \alpha_j \cdot \mathbb{1}[D = j]) \tag{31}$$

where $\Phi(.)$ is the standard normal c.d.f.

In Figure D.1 we report the correlation between willingness-to-pay in 2007 and displacement outcomes in t+1, t+2,... up to t+5. We report for each year the semi elasticity

$$\text{Semi}^{t+k}_{M(\mathbf{p})} = \frac{E(\pi_{t+k}|Z, D = M) - E(\pi_{t+k}|Z, D = 0)}{E(\pi_{t+k}|Z, D = 0)}$$

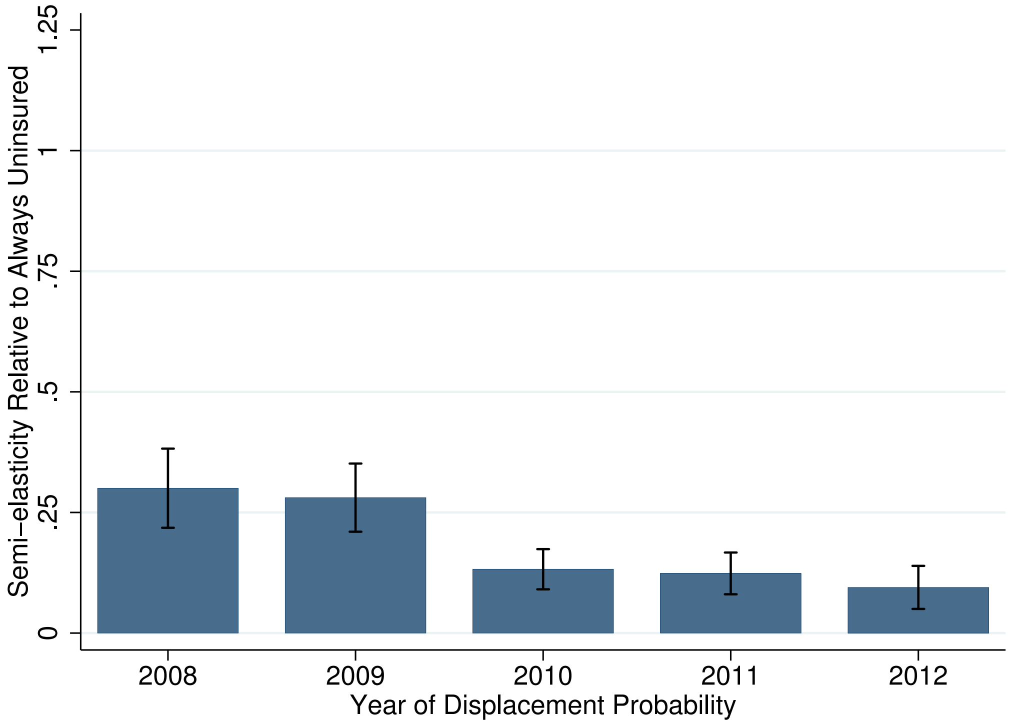
of the displacement probability in $t + k$ for the marginals $D = M$ relative to the individuals in the basic coverage throughout $D = 0$. The figure reveals that the correlation between unemployment risk and willingness-to-pay decreases rapidly as we consider later years, but remains statistically significant up to five years.

Role of Unpriced Heterogeneity The 2007 price reform allows us to investigate how much of the risk-based selection is driven by selection on specific unpriced observables correlated with risk.

We do so by sequentially including in specification (16) a set of controls X: dummies for age, gender, marital status, education (four categories), industry (1-digit code) and and wealth level (quartiles). We then report for each specification the semi-elasticity $\text{Semi}^{X}_{M(\mathbf{p})} = \frac{E(\pi|Z,X,D=M)-E(\pi|Z,X,D=0)}{E(\pi|Z,X,D=0)}$.

Interestingly, the semi-elasticity increases compared to our baseline when including age as a control. Age is therefore a driver of advantageous selection into UI. Adding rich sets of controls

Figure D.1: Price Variation: Using Displacement Probability as a Risk Outcome



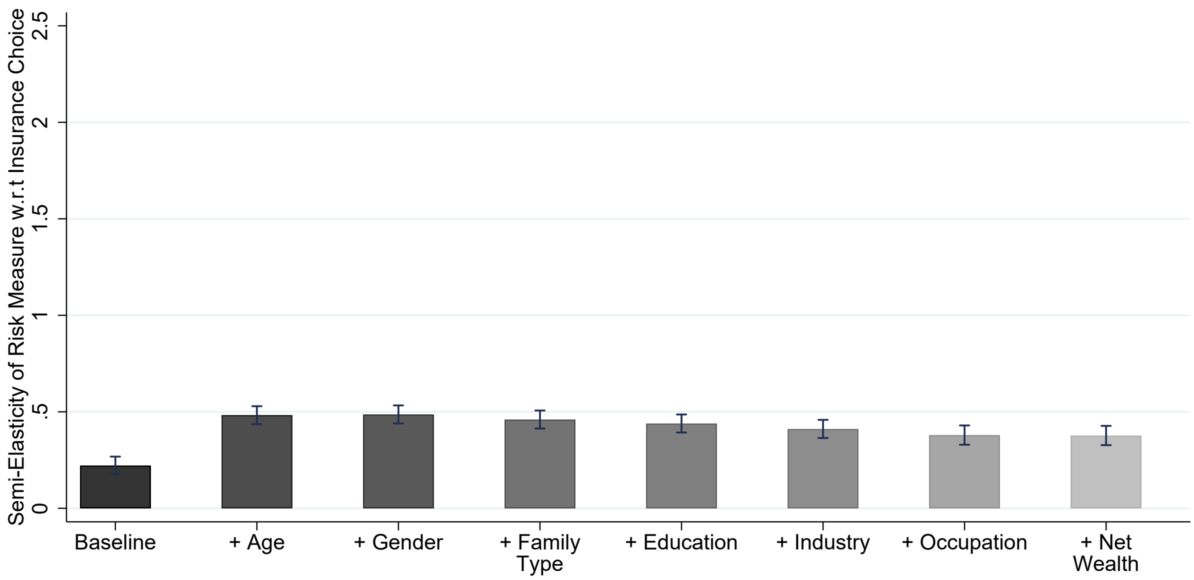
Notes: The Figure reports the correlation between willingness-to-pay in 2007 and realized displacement outcomes in 2008, 2009,.. up to 2012. We report for each year, the semi-elasticity $\text{Semi}^{t+k}_{M(\mathbf{p})}$ of the displacement probability in year $t+k$ for the marginals M relative to the individuals from group 0.

for education, industry, occupation and wealth decreases the estimated correlation only slightly, indicating that there is little risk-related selection along these margins.

Overall, this suggests that demographic characteristics, and age especially, provide advantageous selection on average, such that if contracts were differentiated along these observable dimensions, adverse selection into comprehensive coverage would actually be more severe.

Furthermore, controlling for these unpriced observables does not exhaust risk-based selection in the supplemental UI coverage. In other words, even if the supplemental coverage policy were to price this rich set of observable characteristics, a significant amount of adverse selection would remain.

Figure D.2: Price Variation: Role of Unpriced Heterogeneity

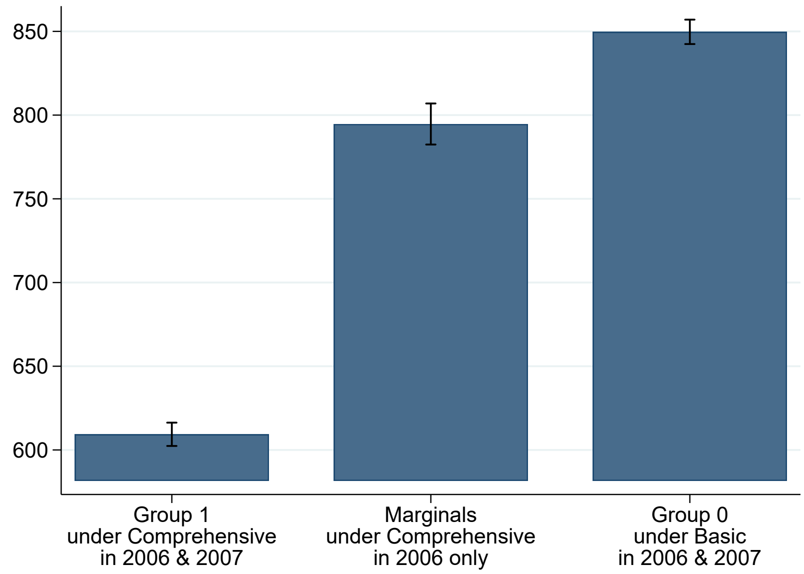


Notes: The Figure explores to what extent estimated adverse selection using the 2007 price variation is driven by selection on observable characteristics that are unpriced in the Swedish UI system. We report the semi-elasticity $\text{Semi}^X_{M(\mathbf{p})}$ of the number of days spent unemployed in 2008 for the marginals M relative to the individuals from group 0. We start with the baseline estimate only controlling for the characteristics affecting the actual UI policy, and show how the semi-elasticity evolves as we add sequentially more characteristics to the vector of controls X. We start by adding demographic controls (age, then gender, and marital status), then controls for skills and other labor market characteristics (controls for education (four categories), industry (1-digit code), occupation (1-digit code) and wealth level (quartiles).

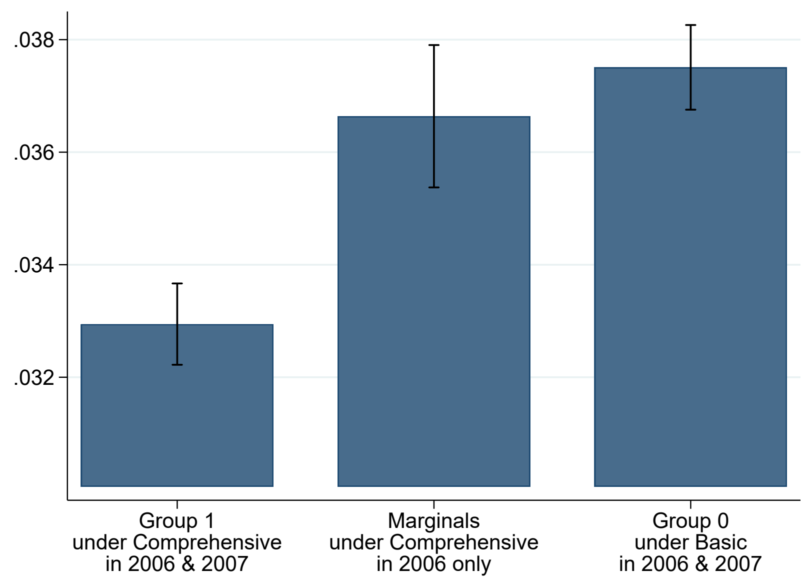
Selection on preferences and value of UI The 2007 price reform also allows to investigate patterns of selection along dimensions other than risk. In Figure D.3, we examine how characteristics that determine the value of unemployment insurance and proxy for risk preferences correlate with willingness-to-pay for insurance revealed by the 2007 price variation. Panel A correlates the level of individuals' net wealth in 2006 in thousands of SEK with their willingness-to-pay controlling for age. Individuals with larger net wealth have more means to smooth consumption in case of displacement, and as a result, should value extra coverage less. The graph indeed confirms the presence of a clear monotonic relationship between net wealth and willingness-to-pay: individuals from group 0 have significantly larger net wealth than the marginals M, who have significantly more net wealth than the individuals from group 1. In panel B, we probe into the potential amount of selection based on risk-preferences. To proxy for risk aversion, we use the fraction of total net wealth invested in risky assets (stocks). The graph shows that the individuals in comprehensive coverage have a significantly larger fraction of risky assets in their portfolio than the marginals and the individuals in basic coverage, conditional on net wealth. This evidence is in line with more risk-averse individuals valuing the extra coverage more.

Figure D.3: Price Variation: Selection on Preferences

A. Net Wealth in 2006 (thousands of SEK)



B. Fraction of Risky Assets in Total Net Wealth

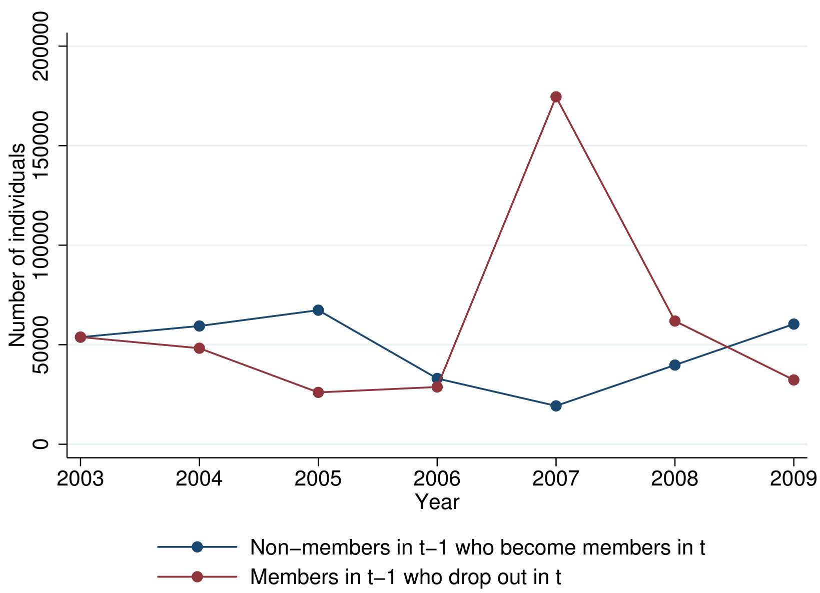


Notes: The Figure uses the 2007 price reform to rank individuals according to their willingness-to-pay for the supplemental coverage $\mathbf{u}$, and uses this ranking to correlate $\mathbf{u}$ with proxies for the value of unemployment insurance and risk preferences. In both panels, individuals are ranked by decreasing order of $\mathbf{u}$. Group 1 on the left are individuals who are insured with the comprehensive coverage both in 2006 and 2007 and have the highest level of $\mathbf{u}$. The middle group corresponds to the marginals ($M(\mathbf{p})$): individuals who were insured with the comprehensive coverage in 2006 but switch out in 2007 when the premium increases. They have a lower level of $\mathbf{u}$ than group 1, but a higher level of $\mathbf{u}$ than the last group on the right (0), of individuals who neither buy the supplemental coverage in 2006, nor in 2007. Using this ranking, we correlate in panel A willingness-to-pay with the level of net wealth in 2006. Individuals with higher net wealth have better means to smooth consumption in case of displacement and should have a lower valuation of additional unemployment insurance. We winsorize net wealth and eliminate the bottom and top percentile of the distribution. In panel B, we proxy for risk aversion using the fraction of total net wealth invested in risky assets (stocks). In both panels we report the average outcome of each group controlling for our baseline vector of characteristics Z plus a cubic polynomial for age, and a cubic for net wealth in panel B.

Robustness Our partition of the population in terms of willingness-to-pay implicitly assumes that $\mathbf{u}$ is constant over time, or to be more precise that the ranking of individuals' willingness-to-pay is the same in 2006 and 2007. In practice $\mathbf{u}$ may change over time, due for instance to idiosyncratic shocks to risk, or preferences, thus creating a flow of individuals switching out of the comprehensive plan, even absent price changes. Appendix Figure D.4 provides evidence that the flow of individuals who switch out of the supplemental coverage was in fact very small prior to the 2007 price reform, but experienced a sudden surge in 2007. This alleviates the concern that our ranking of individuals by willingness-to-pay is confounded by underlying changes in individuals' preferences or risks.

We also note that our partition of the population ignores a negligible fourth group of individuals, who were not buying the comprehensive plan in 2006, but switched in the comprehensive plan in 2007. The size of this group is seven times smaller than the group of individuals switching out of the comprehensive plan in 2007. The ranking of this fourth group in terms of willingness-to-pay is also ambiguous, as one would need to include idiosyncratic shocks to $\mathbf{u}$ to account for the fact that these individuals switched in the comprehensive coverage in 2007 despite the increase in prices $\mathbf{p}$. We display in Appendix Figure D.4 the evolution of the flow of individuals not buying the comprehensive plan in $t-1$ but switching in the comprehensive plan in t. The graph shows that this flow of individuals was small prior to 2007, and equivalent in size to the flow of individuals switching out, hence the stability in the fraction of individuals insured. The flow of individuals switching in seems to decrease with the 2007 reform, but only slightly. The average unemployment risk of the workers switching into the comprehensive plan was the highest among the four groups throughout this period.

Figure D.4: The 2007 Price Reform: Flows of Individuals Switching in and Switching out of the Comprehensive Coverage over Time



Notes: The Figure reports the evolution of the absolute flows of individuals "switching in" and "switching out" of the comprehensive coverage over time. The sample is restricted to individuals were meeting the work eligibility requirement. Individuals who switch in are individuals who were not buying the comprehensive coverage in year $t-1$ but are buying in year t (blue curve). Individuals who switch out are individuals who were buying the comprehensive coverage in year $t-1$ but are no longer buying in year t (red curve). The Figure shows a large and sudden increase in the flow of individuals switching out and a decrease in the flow of individuals switching in, following the large increase in the the premia paid for the supplemental coverage in 2007.

Inertia Inertia is a potentially important behavioral friction, which has been shown to be extremely relevant in other social insurance contexts, such as health insurance. We investigate here the role of inertia and how it affects adverse selection identified in the context of the 2007 price variation. In line with the existing the literature (e.g., Handel (2013)), we use job switchers to proxy for differential exposure to inertia. New employees in a firm arguably face a more active choice environment than existing employees. The former have to reoptimize many choices, while the latter remain in a more passive choice environment. In practice, Figure C.7 above confirms that switching job is indeed associated with a significant change in insurance choices.

In Figure D.5 below, we start by looking at how the price reform of 2007 affected insurance choices for individuals in active choice environments (job switchers) relative to individuals in passive choice environments (job stayers). We find that the 2007 price reform immediately decreased the demand for the comprehensive coverage, in similar proportions, in both the active and the more passive choice environment. But we do find a larger response one year after (in 2008) for job switchers than for non-switchers, which suggests the presence of inertia. Overall, though, the graph suggest that inertia plays a relatively limited role in our setting: individuals in passive choice environments reacted strongly to the reform, and their long run demand response is quite similar to that of individuals facing more active choice environments.

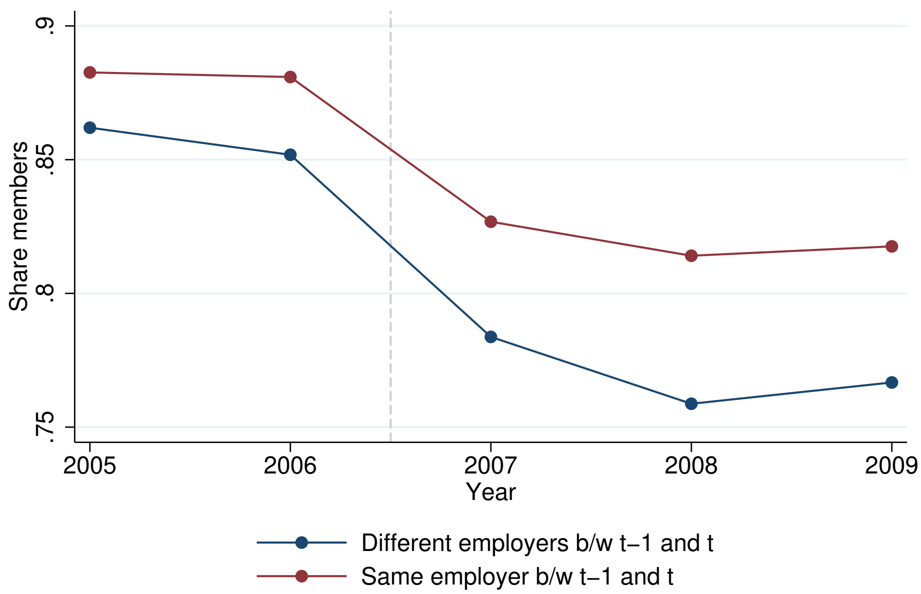
In Figure D.6, we further investigate whether the adverse selection created by these demand responses is different for individuals in active vs passive choice environments. We report the semi-elasticity of the predicted risk $\hat{\pi}_j, j \in \{0, 1\}$ of marginals versus individuals always in the basic coverage, splitting the sample by active vs passive choice environment in 2007:

$$\text{Semi}_{M(\mathbf{p})}^{\hat{\pi}_j} = \frac{E(\hat{\pi}_j|Z, D = M) - E(\hat{\pi}_j|Z, D = 0)}{E(\hat{\pi}_j|Z, D = 0)}$$

where Z is a vector of characteristics affecting the contract space. We find that the adverse selection identified by the 2007 price reform is slightly larger for predicted risk in the basic coverage for workers observed in active compared to workers observed in passive choice environment. But we do not find any significant difference in adverse selection for risk in the comprehensive coverage.

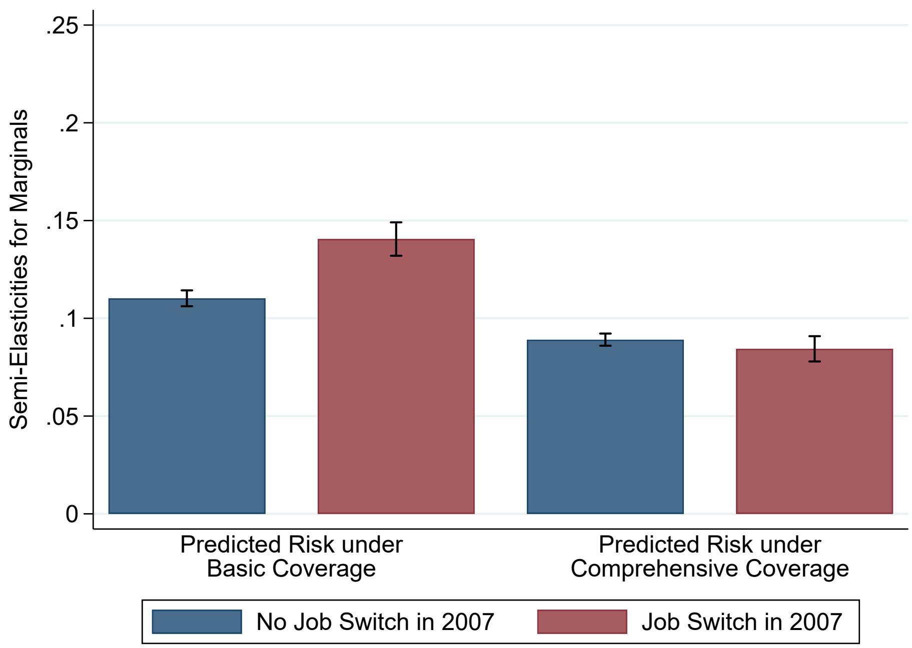
Put together, this evidence suggests that inertia does not seem to critically affect our estimates of the demand and marginal cost curves. It is worth noting though that the relatively modest inertia we find is likely due to the fact that the 2007 reform was large, and salient.

Figure D.5: Inertia: Fraction of Workers Buying the Comprehensive Coverage around the 2007 Reform by Job Switching Status



Notes: The Figure reports the evolution of the fraction of individuals buying the comprehensive coverage around the 2007 by job switching status. In line with the existing the literature (e.g. Handel (2013)), we use job switchers to proxy for differential exposure to inertia. New employees in a firm face a more active choice environment than existing employees. The former have to reoptimize many choices, while the latter remain in a more passive choice environment.

Figure D.6: Inertia & Adverse Selection: Relative Risk of the Marginals Compared to Individuals in the Basic Coverage by Job Switching Status



Notes: The Figure reports the estimated adverse selection created by the 2007 price reform for two sets of workers who are differently exposed to inertia. The red bars refer to individuals who switch job in 2007. These individuals are facing an active choice environment in 2007, at the moment of the price change. The blue bars refer to individuals who stay with their employers. These individuals are facing a passive choice environment in 2007. For both groups of workers, we report the semi-elasticity of the predicted risk $\hat{\pi}_j, j \in \{0, 1\}$ of marginals versus individuals always in the basic coverage:

$$\text{Semi}^{\hat{\pi}_j}_{M(\mathbf{p})} = \frac{E(\hat{\pi}_j|Z, D = M) - E(\hat{\pi}_j|Z, D = 0)}{E(\hat{\pi}_j|Z, D = 0)}$$

where Z is a vector of characteristics affecting the contract space.

Appendix E Benefit Variation: Additional Material

In this appendix we provide additional material regarding our RKD estimation of the effect of benefit variation on insurance choices and risk-based selection:

(i) we present results assessing the validity of our RK design.

(ii) we present results assessing the sensitivity of our RKD estimates.

Table E.2 provides the summary statistics for the sample used for the RKD analysis.

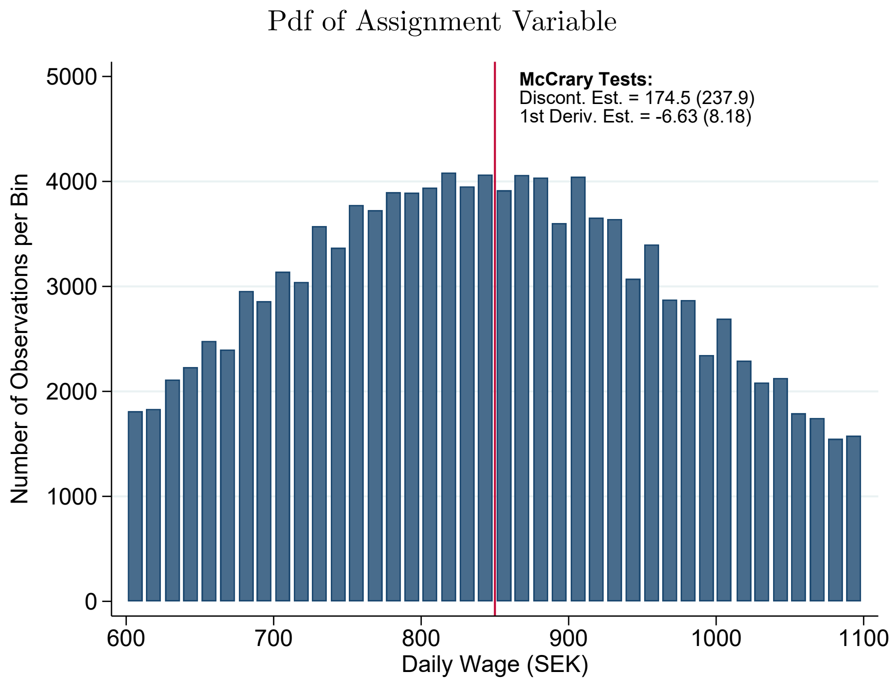
E.1 Assessing Validity of the RK Design

The key identifying assumption of the RK design is the existence of a smooth relationship at the threshold $w = 850SEK$ between the assignment variable and any pre-determined characteristics affecting the demand for insurance. To assess the credibility of this assumption, we conduct two types of analysis [see also Kolsrud et al. [2018]].

Smoothness of the distribution of the assignment variable at the kink First, we focus on the probability density function of the assignment variable, to detect manipulation or lack of smoothness around the kink that could indicate the presence of selection. Figure E.1 shows that the pdf of daily wage does not exhibit a discontinuity nor lack of smoothness at the kink, which is confirmed by the results of formal McCrary tests.

Covariate Tests Second, we investigate the presence of potential selection along observable characteristics around the kink. For this purpose, instead of looking at each characteristics in isolation, we aggregate them in a covariate index. The index is a linear combination of a vector of characteristics X that correlate with our outcomes of interest for the RKD, which includes age, gender, level of education, region, family type and industry. The coefficients in the linear combination are obtained from a regression of the outcome variable on these covariates. In Figure E.2, we display the relationship between this covariate index and the assignment variable for our three outcomes of interest: the choice of coverage, and the predicted risk under basic and comprehensive coverage. The relationship between the index and daily wage appears smooth around the 850SEK threshold. Yet, formal tests of non-linearity suggest the presence of a significant (although economically small) kink at the threshold for insurance choice. But for predicted risk, we do not find any significant non-linearity in the covariate index at the kink.

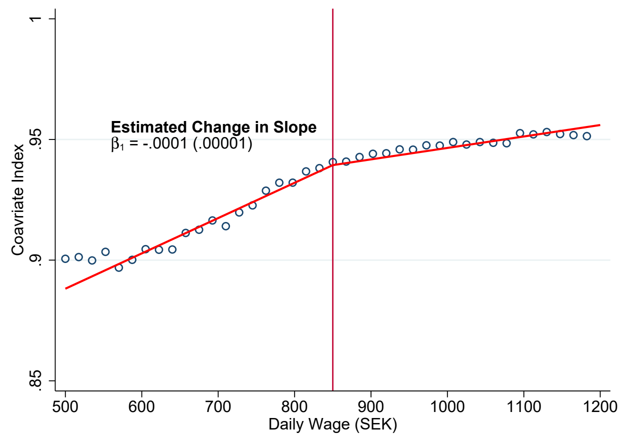
Figure E.1: Regression Kink Design: Testing for Manipulation of Assignment Variable



Notes: The panel displays the probability density function of daily wage. We also report on the graph formal McCrary tests for the existence of a discontinuity and of lack of smoothness of the pdf at the 850SEK threshold.

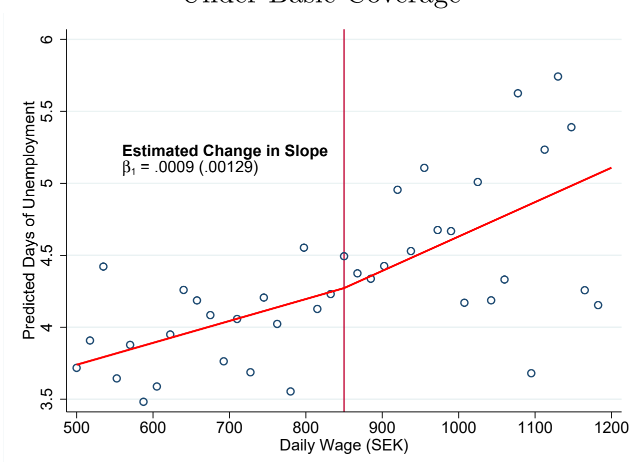
Figure E.2: Regression Kink Design: Smoothness of Distribution of Observables Characteristics

A. Covariate Index vs Assignment Variable: Insurance Choices

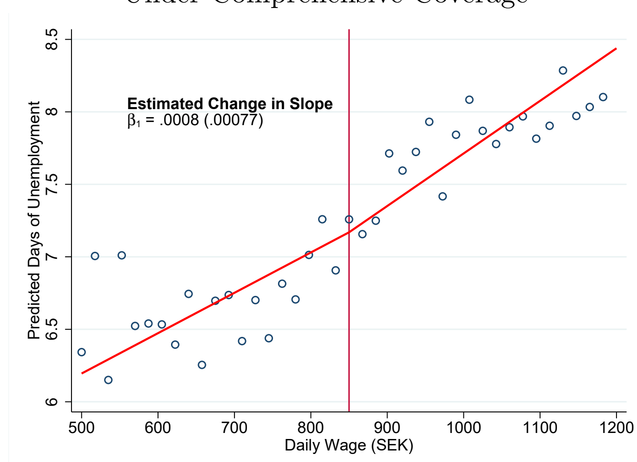


B. Covariate Index vs Assignment Variable: Predicted Risk

Under Basic Coverage



Under Comprehensive Coverage



Notes: The Figure investigates the presence of potential selection along observable characteristics around the kink. For this purpose, we aggregate observable characteristics into a covariate index. The index is a linear combination of a vector of characteristics X that correlate with the outcome, and which includes age, gender, level of education, region, family type and industry. The coefficients in the linear combination are obtained from a regression of the outcome variable on these covariates. Panel A displays the relationship between the assignment variable and the covariate index for the probability to buy the comprehensive coverage. Panel B displays the corresponding graph for the covariate indexes of predicted risk under basic and under comprehensive coverage. We also report on each graph formal tests of non-linearity, i.e. the coefficients β_1 obtained from a specification similar to (18).

E.2 Assessing Sensitivity of the RKD estimates

Sensitivity to bandwidth choice Our baseline RK results use a bandwidth of 350SEK for the daily wage. We start by investigating how sensitive our results are to different bandwidth choices. In Figure E.3, we plot for our three outcomes the value of the RK estimate and its 95% confidence interval for various values of the bandwidth. The graph shows that estimates are stable across bandwidth size. We also computed the optimal bandwidth from Calonico, Cattaneo, & Titiunik (2014), and find 358SEK for the predicted risk and 175SEK for insurance choice.

Sensitivity to inclusion of controls We next investigate how sensitive our results are to the inclusion of the set of controls X. In table E.1, we report in panel A column (1) the estimate of the change in slope β_1 from specification (18), where we do not include the vector X. In column (2), we add controls for age, gender and family types. In column (3), we also add controls for education, region of residence, and industry. We find that the results are stable across these specifications. We then replicate this analysis for predicted risk. In panel B, we focus on predicted risk under basic coverage, and in panel C on predicted risk under comprehensive coverage. Each column reports the estimate β_1 from specification (19), and we vary across columns the set of controls included in the residualization procedure

$$E(\hat{\pi}_j|Z, X) = (1 - f(0|Z, X)) \exp(Z'\gamma_Z + X'\gamma_X)$$

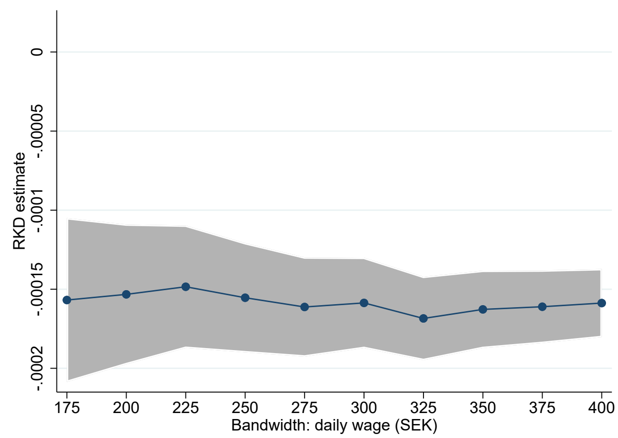
We find that results are also stable to the inclusion of controls.

Inference Finally, we explore the robustness of our inference approach to non-linearities in the relationship between the assignment variable and the outcome. We implement a permutation test and compare the coefficient estimate at the true kink to those at "placebo" kinks placed away from the true kink.

In Figure E.4, we report the probability density function of the estimated change in slope β_1 for 1000 placebo kinks outside the 750SEK-950SEK range. Panel A shows the distribution of placebo RK estimates, using specification (18), for the probability of buying the comprehensive coverage. The estimated coefficient at the true kink lies markedly below all the placebo estimates, indicating that our estimates are unlikely to pick up some non-linearity in the relationship between daily wage and insurance choice. In Panel B we report the distribution of placebo RK estimates for the predicted risk under basic and comprehensive using specification (19). In both cases, we find that the vast majority of placebo estimates is negative, so that if anything, there is non-linearity in the opposite direction than the one detected at the true kink. The probability to find a placebo estimate larger than the estimate at the true kink is, in both cases, inferior to 5%.

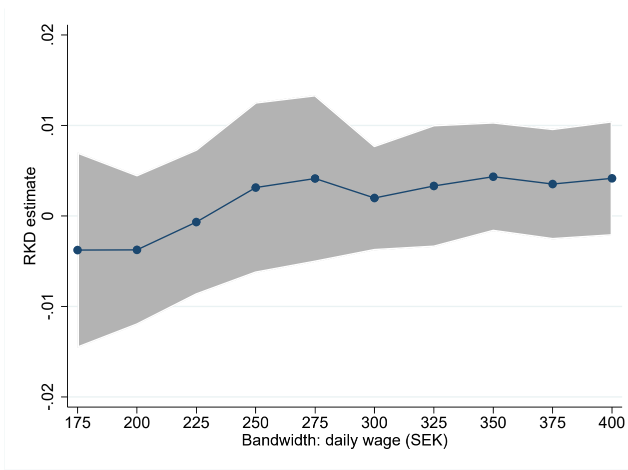
Figure E.3: Regression Kink Design: Sensitivity to Bandwidth Choice

A. RKD Estimates of Insurance Choice Response by Bandwidth

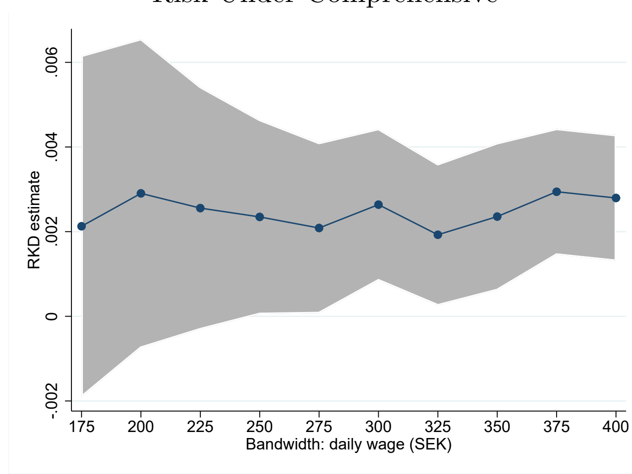


B. RKD Estimates of Predictable Risk by Bandwidth

Risk Under Basic



Risk Under Comprehensive



Notes: The Figure investigates the sensitivity of our estimates to the choice of bandwidth for the RK estimation. Our baseline bandwidth is 350. Panel A plots the value of the RK estimate and its 95% confidence interval for various values of the bandwidth. The graph shows that estimates are stable across bandwidth size. Panel B plots similar graphs for the predicted risk under basic and under comprehensive coverage.

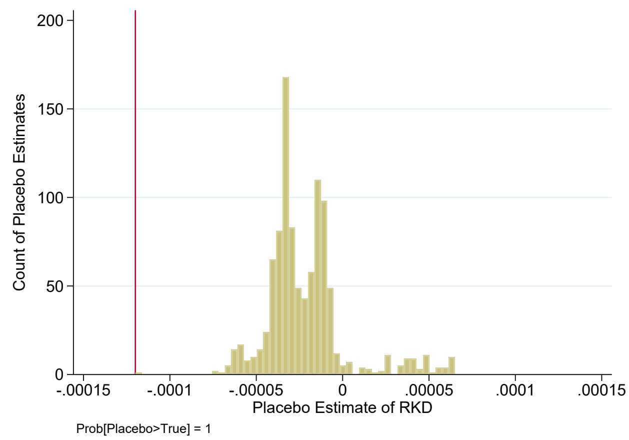
Table E.1: Regression Kink Design: Sensitivity to Inclusion of Controls

	A. Probability to Buy Comprehensive			**B. Risk Under Comprehensive**			**C. Risk Under Basic**		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
β_1	- .016	- .013	- .012	.307	.359	.279	.646	.520	.204
	(.001)	(.001)	(.001)	(.094)	(.093)	(.105)	(.467)	(.330)	(.144)
N	110,123	110,123	110,123	89,576	89,576	89,576	3,998	3,998	3,998
Baseline	×	×	×	×	×	×	×	×	×
Age, gender family type		×	×		×	×		×	×
Education, region industry			×			×			×

Notes: The baseline controls refer to the vector Z of characteristics affecting premia. It consists in a dummy for union membership, a dummy for eligibility and year fixed effects.

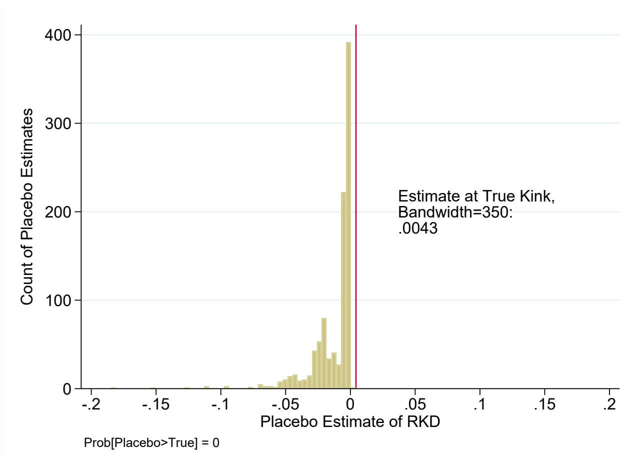
Figure E.4: Regression Kink Design: Permutation-Based Inference

A. Distribution of Placebo Estimates of Insurance Choice Response

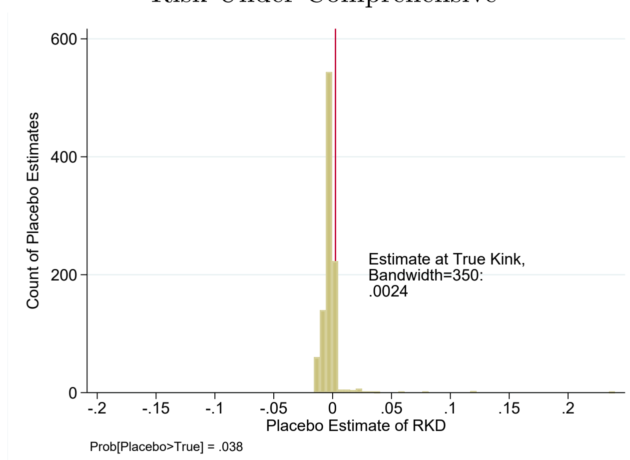


A. Distribution of Placebo Estimates of Predicted Risk

Risk Under Basic



Risk Under Comprehensive



Notes: The Figure reports the probability density function of the estimated change in slope β_1 for 1000 placebo kinks outside the 750SEK-950SEK range. Panel A shows the distribution of placebo RK estimates for the probability of buying the comprehensive coverage. Panel B reports the distribution of placebo RK estimates for the predicted risk under basic and comprehensive. We also report on all three graphs the probability to find a placebo estimate larger than the estimate at the true kink.

Table E.2: Summary Statistics - RKD Population

	Mean	P10	P50	P90
	I. Unemployment			
Days unemployed	180.07	21	145	365
Predicted days unemployed under Basic	4.81	2.33	3.19	5.87
Predicted days unemployed under Comprehensive	8.15	3.88	5.26	11.5
Unemployment spell duration (days)	410	91	301	910
Fraction receiving layoff notification	.04	-	-	-
	II. Union and UI Fund Membership			
Union membership	.78	-	-	-
UI fund membership	.96	-	-	-
Switch from coverage 0 to 1	.04	-	-	-
Switch from coverage 1 to 0	.01	-	-	-
	III. Demographics			
Age	37.95	23	37	55
Years of education	12.18	10	12	16
Fraction men	.56	-	-	-
Fraction married	.33	-	-	-
	IV. Income and Wealth SEK 2003(K)			
Gross earnings	127.6	0	127	246
Net wealth	153.7	-182	0	644
Bank holdings	28.9	0	0	72
N	140,777			

Notes: The Table provides summary statistics for the RKD sample. See Table 1 for our main sample of interest.

Appendix F Welfare Analysis: Additional Material

This appendix provides further detail and derivations for the theoretical analysis in Section 3 and supporting material for the empirical implementation in Section 6.

F.1 Conceptual Framework: Further Details

To evaluate the design of the social insurance system allowing for choice, we characterized the welfare impact of small changes in prices and coverages with two plan options, $\{(b_j, p_j)\}_{j=1,0}$. This relies on the social welfare function being concave and differentiable so that we can characterize the optimal contract using first-order conditions. We also assumed that workers' preferences are quasi-linear in prices so that an individual's risk π_j, conditional on plan choice j, does not depend on prices and neither does the ranking of individuals' valuations $\mathbf{u}_i$. This is a standard assumption in the insurance literature, but the implications from relaxing it are evident from the analysis of coverage changes.

Regarding the timing of the model, we stick closely to the structure of the Swedish UI system where individuals become eligible to receive the supplemental benefits when they have been contributing for one year to the comprehensive coverage, and can opt in and out of the comprehensive plan at any time. As a consequence, the value u_j^t of plan j in year t depends on unemployment risk π^{t+1}, the expected number of days spent unemployed in year $t+1$. With this in mind, we have dropped time subscripts, but $u\left(\pi\right)$ always refers to $u^t\left(\pi^{t+1}\right)$, unless otherwise specified.

Our framework is highly stylized, but allows for multi-dimensional heterogeneity and endogenous actions. Following a recent tradition in the social insurance literature [see Chetty and Finkelstein [2013]], we choose not to explicitly model the underlying heterogeneity and actions. The key microfoundations for our analysis are the resulting plan valuations and costs and how they change with the plans' prices and coverage levels. For example, in a setting with expected utility and binary unemployment risk, the value of a plan to a worker equals

$$u_i\left(b_j, p_j\right) = \max_{a'} \pi\left(a'|\theta_i\right) \tilde{u}\left(b_j - p_j, a'|\mu_i\right) + \left(1 - \pi\left(a'|\theta_i\right)\right) \tilde{u}\left(w_i - p_j, a'|\mu_i\right), \tag{32}$$

while the insurer's cost equals $c_i\left(b_j, p_j\right) = \pi(a'|\theta_i)b_j$. The value and cost are interdependent through the risk parameter θ and the effort choice a, which in turn depends on the preference parameter μ that also affects the valuation. In general, a worker's expected utility depends both on the probability of job loss and the time spent unemployed. The government's expected cost would depend only on the expected number of days spent unemployed when the benefit profile is flat.

Sorting Effect The fiscal externality in both Propositions 1 and 2 depends on how many individuals change in response to the policy (as captured by the demand elasticity) and the cost characteristics of those who switch. We develop here formally why the cost characteristics of the switchers in response to a change in coverage are different than for a change in price under multi-dimensional heterogeneity.

We first re-write the social welfare function, ranking individuals based on their utility gain of the comprehensive relative to the basic plan,

$$\begin{aligned}\mathcal{W} &\equiv \int_{\mathbf{u}_i\geq 0}\omega\left(u_i\left(b_1,p_1\right)\right)di+\int_{\mathbf{u}_i<0}\omega\left(u_i\left(b_0,p_0\right)\right)di+\lambda\left\{F_1\left[p_1-E_1\left(\pi_1\right)b_1\right]+F_0\left[p_0-E_0\left(\pi_0\right)b_0\right]\right\},\\ &= \int_{\mathbf{u}\geq 0}E(\omega\left(u\left(b_1,p_1\right)\right)|\mathbf{u})dG\left(\mathbf{u}\right)+\int_{\mathbf{u}<0}E(\omega\left(u\left(b_0,p_0\right)\right)|\mathbf{u})dG\left(\mathbf{u}\right)\\ &\quad +\lambda\left\{\left(1-G\left(0\right)\right)\left[p_1-E_1\left(\pi_1\right)b_1\right]+G\left(0\right)\left[p_0-E_0\left(\pi_0\right)b_0\right]\right\}.\end{aligned}$$

Here $G(\cdot)$ is the distribution of $\mathbf{u} = u_1 - u_0$, which depends on the plan characteristics, with $G(0) = F_0$ and $1 - G(0) = F_1$. Following the derivation in Veiga and Weyl [2016] and Handel et al. [2019], we then find

$$\frac{\partial}{\partial x_j}\left[\left(1-G\left(0\right)\right)E_1\left(z_1\right)\right]=E\left(z_1\frac{\frac{\partial \mathbf{u}}{\partial x_j}}{E\left(\frac{\partial \mathbf{u}}{\partial x_j}|\mathbf{u}=0\right)}|\mathbf{u}=0\right)\frac{\partial\left(1-G\left(0\right)\right)}{\partial x_j}. \tag{33}$$

assuming no direct effect of the policy variable x_j on the outcome z_1.

The argument proceeds as following. First, using iterated expectations and introducing notation $\mathbf{u}' \equiv \frac{\partial \mathbf{u}}{\partial x_j}$, we can write

$$\begin{aligned}\frac{\partial}{\partial x_j}\left[\left(1-G\left(0\right)\right)E_1\left(z_1\right)\right] &= \frac{\partial}{\partial x_j}\left[\int_{\mathbf{u}\geq 0}E\left(z_1|\mathbf{u}\right)dG\left(\mathbf{u}\right)\right]\\ &= \frac{\partial}{\partial x_j}\left[\int_{\mathbf{u}\geq 0}\int E\left(z_1|\mathbf{u},\mathbf{u}'\right)f\left(\mathbf{u}'|\mathbf{u}\right)g\left(\mathbf{u}\right)d\mathbf{u}'d\mathbf{u}\right]\\ &= \int\frac{\partial}{\partial x_j}\left[\int_{\mathbf{u}_\varepsilon\geq -\mathbf{u}'\times[x_j-x_\varepsilon]}E\left(z_1|\mathbf{u}_\varepsilon,\mathbf{u}'\right)g^\varepsilon\left(\mathbf{u}_\varepsilon|\mathbf{u}'\right)d\mathbf{u}_\varepsilon\right]f\left(\mathbf{u}'\right)d\mathbf{u}'\end{aligned}$$

The last equality follows from (1) using $f\left(\mathbf{u}'|\mathbf{u}\right)g\left(\mathbf{u}\right) = g\left(\mathbf{u}|\mathbf{u}'\right)f\left(\mathbf{u}'\right)$, (2) approximating $\mathbf{u}\left(b_j\right) \cong \mathbf{u}\left(b_\varepsilon\right) + \mathbf{u}' \times \left[b_j - b_\varepsilon\right]$**,** and substituting the variable in the integral $\mathbf{u}\left(b_j\right)(\equiv \mathbf{u})$ by $\mathbf{u}\left(b_\varepsilon\right)(\equiv \mathbf{u}_\varepsilon)$, where $d\mathbf{u} = d\mathbf{u}_\varepsilon$, conditional on $\mathbf{u}'$.

We can then apply Leibniz rule and find after re-substituting,

$$\begin{aligned}\frac{\partial}{\partial x_j}\left[\int_{\mathbf{u}\geq 0}E\left(z_1|\mathbf{u}\right)dG\left(\mathbf{u}\right)\right] &= \int\left[E\left(z_1\mathbf{u}'|\mathbf{u}=0,\mathbf{u}'\right)f\left(\mathbf{u}'|\mathbf{u}=0\right)d\mathbf{u}'\right]g\left(0\right)\\ &= E\left(z_1\frac{\partial \mathbf{u}}{\partial x_j}|\mathbf{u}=0\right)g\left(0\right).\\ &= E\left(z_1|\mathbf{u}=0\right)E\left(\frac{\partial \mathbf{u}}{\partial x_j}|\mathbf{u}=0\right)g\left(0\right)+cov\left(z_1,\frac{\partial \mathbf{u}}{\partial x_j}|\mathbf{u}=0\right)\end{aligned}$$

Note also that the effect on the share of individuals buying comprehensive coverage equals

$$\frac{\partial\left[1-G\left(0\right)\right]}{\partial x_{j}}=\frac{\partial}{\partial x_{j}}\left[\int_{\mathbf{u}\geq 0}dG\left(\mathbf{u}\right)\right]=E\left(\frac{\partial\mathbf{u}}{\partial x_{j}}|\mathbf{u}=0\right)g\left(0\right).$$

Taken together, we can then indeed write

$$\frac{\partial}{\partial x_{j}}\left[\left(1-G\left(0\right)\right)E_{1}\left(z_{1}\right)\right]=E\left(z_{1}\frac{\frac{\partial\mathbf{u}}{\partial x_{j}}}{E\left(\frac{\partial\mathbf{u}}{\partial x_{j}}|\mathbf{u}=0\right)}|\mathbf{u}=0\right)\frac{\partial\left(1-G\right)}{\partial x_{j}}.$$

Similarly, we can find

$$\frac{\partial}{\partial x_{j}}\left[G\left(0\right)E_{0}\left(z_{0}\right)\right]=E\left(z_{0}\frac{\frac{\partial\mathbf{u}}{\partial x_{j}}}{E\left(\frac{\partial\mathbf{u}}{\partial x_{j}}|\mathbf{u}=0\right)}|\mathbf{u}=0\right)\frac{\partial G}{\partial x_{j}}$$

Applying this now to the difference in average risk for switchers in response to a coverage and price change, we find

$$E_{M(b_{j})}(\pi_{k})-E_{M(\mathbf{p})}(\pi_{k})=\frac{cov\left(\pi_{k},\frac{\partial u_{j}}{\partial b_{j}}|\mathbf{u}=\mathbf{0}\right)}{E\left(\frac{\partial u_{j}}{\partial b_{j}}|\mathbf{u}=\mathbf{0}\right)}.$$

Individuals with higher unemployment risk tend to value extra coverage more. However, with heterogeneity in risk aversion (in addition to risk heterogeneity), the correlation between risks and the marginal value of coverage among the marginal buyers can become negative [Ericson et al. [Forthcoming]]. Relatedly, the risk-based selection into supplemental coverage may well worsen as the basic coverage level increases. The reason is that the variation in willingness-to-pay coming from heterogeneity in risk aversion, which would mute the risk-based selection, decreases as the basic coverage becomes more generous [see Ericson et al. [Forthcoming]].

F.2 Proof of Proposition 1

We denote again by $\mathbf{p}$ and $\mathbf{c}$, the difference in prices and costs between the two plans, e.g. $\mathbf{p}=p_{1}-p_{0}$.

Proof. Differentiating the social welfare with respect to p_{0}, we get:

$$\begin{aligned}\frac{\partial\mathcal{W}}{\partial p_{0}}&=G\left(0\right)E_{0}\left(\frac{\partial\omega_{0}}{\partial p_{0}}\right)\\&+\lambda\left\{G\left(0\right)-\frac{\partial G\left(0\right)}{\partial p_{0}}\mathbf{p}-\frac{\partial\left[\left(1-G\left(0\right)\right)E_{1}\left(c_{1}\right)+G\left(0\right)E_{0}\left(c_{0}\right)\right]}{\partial p_{0}}\right\}.\end{aligned}$$

Here, we are invoking the envelope theorem for the resorting of marginal individuals, $u_{i}\left(b_{1},p_{1}\right)=$

$u_i(b_0, p_0)$. We can rewrite this expression as

$$\frac{\partial \mathcal{W}}{\partial p_0} = G(0) E_0\left(\frac{\partial \omega_0}{\partial p_0}\right) + \lambda\left\{G(0) - (\mathbf{p} - E_M(\mathbf{c}))\frac{\partial G(0)}{\partial p_0}\right\}$$

using the result in equation (33),

$$\begin{aligned}\frac{\partial\left[(1-G(0))E_1(c_1) + G(0)E_0(c_0)\right]}{\partial p_0} &= E_M(c_1)\frac{\partial[1-G(0)]}{\partial p_0} + E_M(c_0)\frac{\partial G(0)}{\partial p_0} \\ &= -E_M(\mathbf{c})\frac{\partial G(0)}{\partial p_0}.\end{aligned}$$

Note that the quasi-linear assumption implies that prices do not cause any moral hazard response, $E_0\left(\frac{\partial c_0}{\partial p_0}\right) = E_0\left(b_0\frac{\partial \pi_0}{\partial p_0}\right) = 0$. Hence, the only impact on the budget constraint is the re-sorting response. This response itself also simplifies due to the quasi-linearity assumption. Since $\frac{\partial \mathbf{u}}{\partial p_0}$ is constant, it depends on the demand response $\frac{\partial G(0)}{\partial p_0}$ multiplied by the cost of providing the supplemental coverage to workers at the margin, which simplifies to the unweighted average among the marginals, $E_M(\mathbf{c}) = E(\mathbf{c}|\mathbf{u}=0)$. Using $\frac{\partial G(0)}{\partial p_0} = -\frac{\partial G(0)}{\partial \mathbf{p}}$, we then find that the FOC with respect to the price of basic coverage, $\frac{\partial \mathcal{W}}{\partial p_0} = 0$, is equivalent to

$$-E_0\left(\frac{\partial \omega_0}{\partial p_0}\right) = \lambda\left\{1 + AS_{\mathbf{p}}\frac{\partial \ln G(0)}{\partial \mathbf{p}}\right\}.$$

In a similar way, we can get the FOC for the price of the comprehensive plan

$$-E_1\left(\frac{\partial \omega_1}{\partial p_1}\right) = \lambda\left\{1 + AS_{\mathbf{p}}\frac{\partial \ln(1-G(0))}{\partial \mathbf{p}}\right\}$$

Putting the two FOCs together, we get the expression in the Proposition. □

F.3 Proof of Proposition 2

Proof. We consider the welfare impact of an increase in b_0, for given prices and coverage b_1. The impact of a change in b_0 on the government's budget depends both on the change in selection into both plans and the direct effect from increasing the coverage,

$$\frac{\partial}{\partial b_0}\left[\int_{\mathbf{u}\geq 0}[p_1 - E(c_1|\mathbf{u})]\,dG(\mathbf{u}) + \int_{\mathbf{u}<0}[p_0 - E(c_0|\mathbf{u})]\,dG(\mathbf{u})\right]$$

$$\begin{aligned}&= \left[\mathbf{p} - E\left(\mathbf{c}\times\frac{\frac{\partial \mathbf{u}}{\partial b_0}}{E\left(\frac{\partial \mathbf{u}}{\partial b_0}|\mathbf{u}=0\right)}|\mathbf{u}=0\right)\right]\frac{\partial(1-G(0))}{\partial b_0} - G(0)\frac{\partial E_0 c_0}{\partial b_0} \\ &\equiv -\left[\mathbf{p} - E_{M(b_0)}(\mathbf{c})\right]\frac{\partial G(0)}{\partial b_0} - G(0)\frac{\partial E_0 c_0}{\partial b_0}\end{aligned}$$

By analogy to the subsidy change, we decompose the change in cost from providing coverage due to the change in selection as the demand effect $\frac{\partial G}{\partial b_0}$ multiplied by the fiscal externality $\mathbf{p}-E_{M(b_0)}(\mathbf{c})$, caused by the switching of individuals who respond to the coverage change. This fiscal externality differs from the fiscal externality of the subsidy as different individuals respond to a change in coverage depending on their marginal value of basic coverage $\frac{\partial \mathbf{u}}{\partial b_0}$, explaining the weights put on the costs of the different marginals with $\mathbf{u} = 0$. This is discussed in detail in Appendix F.1. In addition to the selection response, an increase in coverage of the basic plan affects the government's expenditures directly, but also indirectly through a moral hazard response. That is,

$$\frac{\partial E_0 c_0}{\partial b_0} = E_0(\pi_0) + E_0\left(\frac{\partial \pi_0}{\partial b_0}\right) b_0 = E_0(\pi_0)\left[1 + FE^{MH}_{b_0}\right].$$

Invoking now the envelope condition for the individuals at the margin (i.e., $\mathbf{u} = 0$), we find

$$d\mathcal{W} = G(0)\, E_0\left(\frac{\partial \omega_0}{\partial b_0}\right) - \lambda G(0)\, E_0(\pi_0)\left[1 + FE^{MH}_{b_0}\right] - \lambda\left[\mathbf{p} - E_{M(b_0)}(\mathbf{c})\right]\frac{\partial G}{\partial b_0},$$

where

$$E_0\left(\frac{\partial \omega_0}{\partial b_0}\right) = \frac{1}{G(0)}\int_{\mathbf{u}<\mathbf{0}} E\left(\omega'(u_0)\frac{\partial u(b_0, p_0)}{\partial b_0}|\mathbf{u}\right) dG(\mathbf{u})$$

Hence, at an (interior) optimum, we need $d\mathcal{W}=0$ and thus

$$E_0\left(\frac{\partial \omega_0}{\partial b_0}\right)/E_0(\pi_0) = \lambda\left\{1 + FE^{MH}_{b_0} + \left[\mathbf{p} - E_{M(b_0)}(\mathbf{c})\right]\frac{\partial \ln G}{\partial b_0}/E_0(\pi_0)\right\}.$$

In a similar way, we can get the FOC for the coverage of the comprehensive plan,

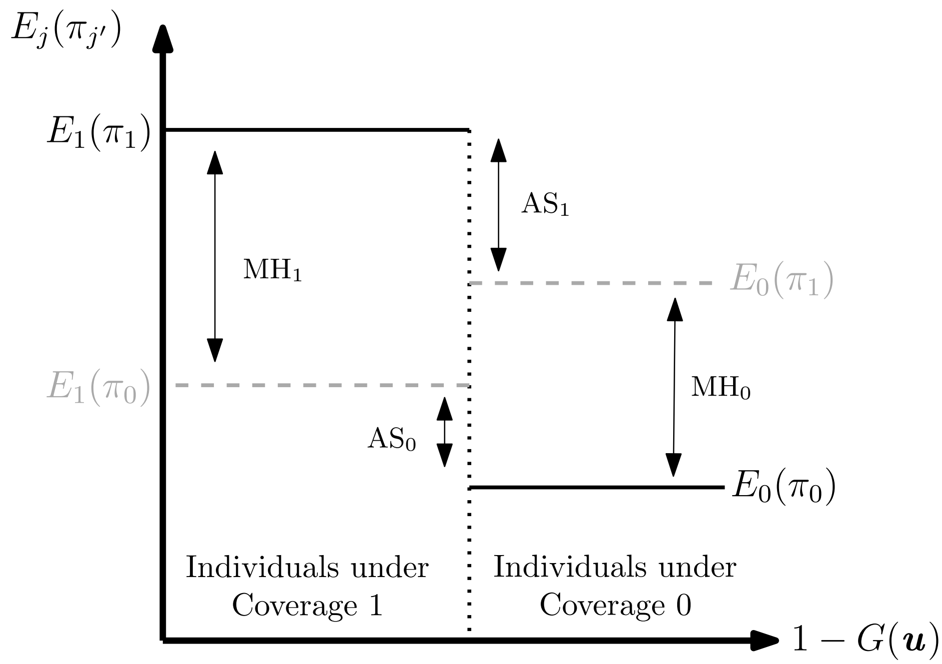
$$E_1\left(\frac{\partial \omega_1}{\partial b_1}\right)/E_1(\pi_1) = \lambda\left\{1 + FE^{MH}_{b_1} - \left[\mathbf{p} - E_{M(b_1)}(\mathbf{c})\right]\frac{\partial \ln(1-G)}{\partial b_1}/E_1(\pi_1)\right\}.$$

Putting the two FOCs together, we get the expression in the Proposition.

□

F.4 Graphical Representation of AS vs. MH

Figure F.1: Decomposition of PCT Statistic



Notes: The Figure illustrates the decomposition of the positive correlation test (PCT) statistic $E_1(\pi_1) - E_0(\pi_0)$ studied in Section 3. Workers opt for the comprehensive plan 1 when $\mathbf{u} \geq 0$ and for the basic plan 0 otherwise.$E_j(\pi_{j'})$ denotes the average unemployment risk for workers who opt for coverage j when under plan j'. There are two complementary ways to quantify the respective roles of adverse selection and moral hazard due to the fact that the measurement of the differences in risk due to adverse selection is plan-dependent, while the measurement of the differences in risk due to moral hazard is group-dependent. A first decomposition consists of adverse selection in the comprehensive plan (AS_1) plus moral hazard for the group selecting basic coverage (MH_0). A second decomposition consists of moral hazard for the group selecting comprehensive coverage (MH_1) and adverse selection in the basic plan (AS_0). Relating this to the textbook analysis of selection and treatment effects, the moral hazard response can be interpreted as the treatment effect on risk from providing supplementary coverage. This treatment effect can be different for workers who select into treatment compared to those who do not. The difference in treatments effects between the two groups depends on the difference in risks under the comprehensive and basic coverage respectively.

F.5 Policy Implications: Supporting Material

Risk and Cost Curves Using the price variation, we can infer how the risk under basic and comprehensive coverage changes with willingness-to-pay for the supplemental coverage. In Section 6, we convert the resulting *marginal* risk curves into marginal cost curves like in Einav et al. [2010b], accounting for the coverage levels. The risk curves are displayed in Figure F.2. Juast like for the cost curves, we position the three groups of individuals $i = 1, M, 0$ on the x-axis according to their willingness-to-pay. Individuals who choose the basic coverage (0) both in 2006 and 2007 are on the right hand side, while individuals who always buy the comprehensive coverage (1) are on the left-hand side of the graph. The marginals correspond to the group in between, and their share is given in Figure 4, where we see the fraction of individuals buying dropping from 86 to 78% with the 2007 price increase. We plot with green dots the observed average realized risk, measured by the number of days spent unemployed in year 2008, for the three groups. Note that all risk measures in Figure F.2 are conditional on the vector of characteristics Z, and normalized to the average risk under basic coverage of individuals observed under basic coverage $E_0(\pi_0)$. The difference in realized risk (green dots) between the marginals and group 0 is therefore equivalent to our semi-elasticity estimate $\text{Semi}_{M(\mathbf{p})}^{Baseline}$.

We then plot with blue triangles the average predicted risk under basic coverage and with red triangles the average predicted risk under comprehensive coverage for all three groups. The blue triangles identify the marginal risk curve in the basic coverage, while the red triangles plot the marginal risk curve in the comprehensive coverage. These two curves provide all the information necessary to compute the fiscal externality term $FE_{\mathbf{p}}^{AS}$ that determines the optimal price level in Proposition 1. Moreover, the difference between the red and blue triangles for each group identifies the moral hazard of moving these individuals from basic to comprehensive coverage.

Consumption Smoothing Gains When considering to further differentiate the coverage levels following Proposition 2, the fiscal externalities need to be traded off against the relative consumption smoothing gains from extra coverages: $\frac{E_1\left(\frac{\partial \omega_1}{\partial b_1}\right)/E_1(\pi_1)}{E_0\left(\frac{\partial \omega_0}{\partial b_0}\right)/E_0(\pi_0)}$. Given our estimates of the fiscal externalities, further differentiation would therefore be socially valuable if the marginal utility of an increase in benefits for individuals under the comprehensive coverage is worth at least 27% more than the marginal utility of an increase in benefits for individuals under the basic coverage.

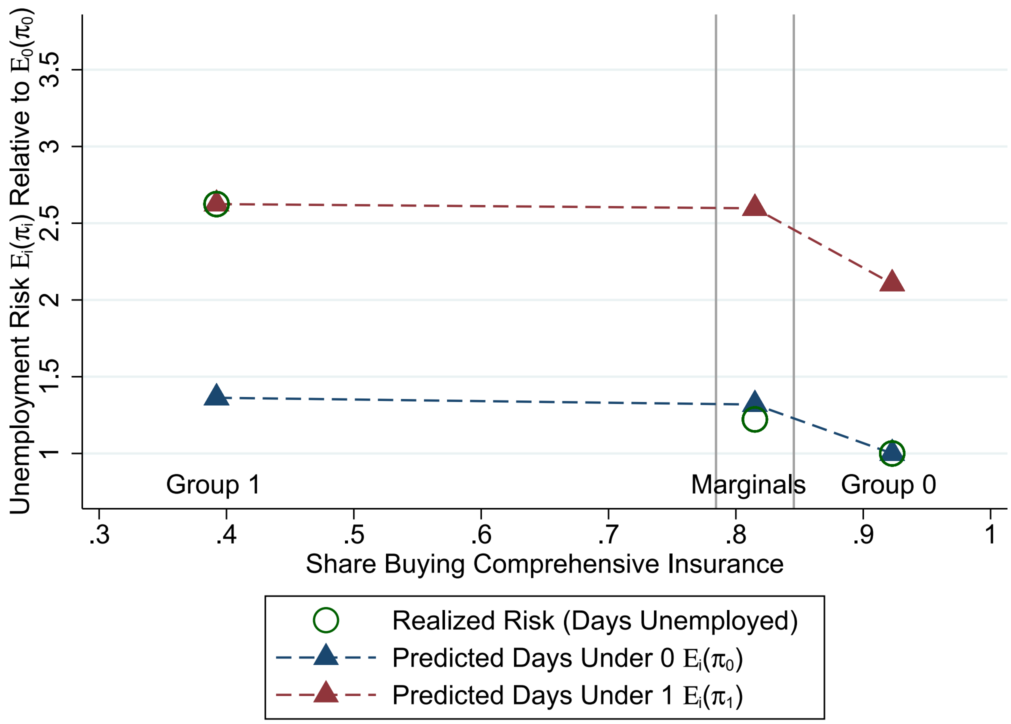
To estimate the magnitude of the relative consumption smoothing gains, we need to know about the two basic forces that underlie them. The first is heterogeneity in the value of insurance, conditional on risk: people who buy the comprehensive coverage may do so because they value an additional kroner of consumption when unemployed more. Landais and Spinnewijn [Forthcoming] find in the Swedish context that the mark-up workers under comprehensive coverage are willing to pay for the supplemental coverage is 160% larger compared to workers under basic coverage, conditional on their risk: $E_1(\overline{MRS}_{b_0,b_1}) = 2.6 \cdot E_0(\overline{MRS}_{b_0,b_1})$.[51] This number underlines that there is indeed significant heterogeneity in willingness-to-pay for insurance conditional on risk, which is a strong force pushing for coverage differentiation.

The second force is diminishing marginal utility of consumption. The marginal utility of further coverage should be estimated at coverage level b_1 for individuals buying the comprehensive coverage and at coverage level b_0 for individuals in the basic plan. Diminishing marginal utility of consumption makes the value of an additional kroner lower when evaluated at b_1 than at b_0.

Using a Taylor approximation, we can provide a back-of-the envelope calculation of the relative consumption smoothing gains from extra coverage. We can indeed write that: $E_j\left(\frac{\partial \omega_j}{\partial b_j}\right)/E_j\left(\pi_j\left(b_j\right)\right) \equiv$

[51] The mark-up $\overline{MRS}_{b_0,b_1}$ is defined as the average marginal rate of substitution between consumption when employed and unemployed evaluated between coverage level b_0 and b_1. These estimates come from the model of Table 3 column (2), for which they find $E_1\left(\overline{MRS}_{b_0,b_1}\right) = 2.92$, $E_0\left(\overline{MRS}_{b_0,b_1}\right) = 1.12$

Figure F.2: Price Variation: Decomposition and Marginal Risk Curves



Notes: The Figure uses estimates of Figure 5 to trace out the marginal risk curves under basic coverage and comprehensive coverage using the sample of all eligible workers observed in 2006 and 2007. We start by positioning the three groups of individuals $j = 1, M, 0$ on the x-axis according to their willingness-to-pay. Individuals who choose the basic coverage (0) both in 2006 and 2007 are on the right hand side, while individuals who always buy the comprehensive coverage (1) are on the left-hand side of the graph. The marginals correspond to the group in between, and their share is given in Figure 4, where we see the fraction of individuals buying dropping from 86 to 78% with the 2007 price increase. We plot with green dots the observed average realized risk, measured by the number of days spent unemployed in year $t+1$, for the three groups. All risk measures are conditional on the vector of characteristics Z, and normalized to the average risk under basic coverage of individuals observed under basic coverage $E_0(\pi_0)$. The difference in realized risk (green dots) between the marginals and group 0 is therefore equivalent to our semi-elasticity estimate $\text{Semi}_{M(\mathbf{p})}^{Baseline}$. We then plot with blue triangles the average predicted risk under basic coverage and with red triangles the average predicted risk under comprehensive coverage for all three groups. The blue triangles identify the marginal risk curve in the basic coverage, while the red triangles plot the marginal risk curve in the comprehensive coverage. The difference between the red and blue triangles for each group identifies the moral hazard of moving these individuals from coverage 0 to coverage 1. These two curves provide all the information necessary to compute the fiscal externality term $FE_{\mathbf{p}}^{AS}$ that determines the optimal price level in Proposition 1.

$E_j \left(u' \left(c_u \left(b_j\right)\right) / u' \left(c_e \left(b_j\right)\right)\right) \cong 1 + \gamma \times E_j \left(\frac{\Delta c}{c} \left(b_j\right)\right)$, where c_u and c_e denote the consumption levels when unemployed and employed, and $\gamma = \frac{u''(c)c}{u'(c)}$ is the parameter of relative risk aversion. Similarly, using a linear approximation for the mark-up, we have that: $E_j \left(\overline{MRS}_{b_0,b_1}\right) \cong 1 + \gamma \times E_j \left(\frac{\Delta c}{c} \left(b_0\right) + \frac{\Delta c}{c} \left(b_1\right)\right)/2$. This allows to link the average mark-up $\overline{MRS}_{b_0,b_1}$ estimated in Landais and Spinnewijn [Forthcoming], to the marginal mark-ups evaluated at b_0 and b_1 respectively. In practice, note that we unfortunately cannot observe the counterfactual consumption wedge between employment and unemployment $\left(\frac{\Delta c}{c} \left(b_j\right)\right)$ for workers when put under a different plan j. However, we can easily provide bounds for these counterfactual drops. Assuming the consumption drop were to double when changing from comprehensive to basic coverage (i.e. $E_j(\frac{\Delta c}{c}(b_1)) = E_j(\frac{\Delta c}{c}(b_0))/2$), we get a value of 1.97 for the left-hand side in Proposition 2. Note that Landais and Spinnewijn [Forthcoming] find that the difference in consumption wedges for workers under comprehensive and basic coverage is actually quite small.[52] Assuming that the consumption drop doubles when changing from comprehensive to basic coverage is therefore a conservative upper bound. In other words, estimates from Landais and Spinnewijn [Forthcoming] suggest that the value of extra coverage is likely much more than 27% larger for workers under comprehensive coverage compared to workers under basic coverage, even when evaluated at b_1 and b_0 respectively. It follows that further differentiation in coverage levels would probably enhance welfare.

[52] They indeed find that $E_0 \frac{\Delta c}{c}(b_0) = -.178(.25)$ and $E_1 \frac{\Delta c}{c}(b_1) = -.138(.036)$.