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FUNDAMENTALLY REFORMING THE DI SYSTEM: EVIDENCE FROM GERMAN NOTCH COHORTS

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ABSTRACT

Since 2001, cohorts born after 1960 are no longer eligible for public “Occupational Disability Insurance (ODI)” in Germany. ODI covers health shocks that prevent employees from working in their previous occupation. However, the affected “notch cohorts” are still insured by regular public DI, which covers work disability in any occupation. First, using administrative and survey data, we show that the reform significantly reduced the inflow of new DI beneficiaries by more than 30% in the long-run. Second, we study interaction effects with the private ODI market. Using representative data, we do not find much evidence that the notch cohorts purchased individual private ODI policies at significantly higher rates to compensate for the reduced generosity of the public DI system. To explain such low take-up, we employ a general equilibrium model featuring the roles of the social safety net, administrative costs, and asymmetric information. These driving forces help explain three stylized facts in the individual experience-rated private market for ODI policies: (1) low private ODI take-up and interaction effects with the public system—despite a high lifecycle work disability risk, (2) strong and positive income and health gradients in private ODI take-up, and (3) inversely related income and health gradients in work disability risk over the lifecycle. Simulations illustrate that policy reforms to lower administrative costs have the greatest potential to foster take-up and flatten its income and health gradients.

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1 Introduction

For decades, the question of how to design "optimal" social insurance systems has been at the core of economic research (Chetty and Saez, 2010; Chetty and Finkelstein, 2013; Luttmer and Samwick, 2018; Goodman-Bacon, 2018; Cabral and Cullen, 2019). While countries around the world have organized their social insurance and safety net systems differently, three integral strands exist in every OECD (2010) country: unemployment insurance (Lalive et al., 2015; Hendren, 2017), Workers' Compensation (Powell and Seabury, 2018) and public disability insurance (Koning and Lindeboom, 2015; Autor et al., 2016). What's more, their design and structure are similar across countries. Consequently, experiences from one OECD country might hold important lessons for others (Burkhauser et al., 2016).

In the United States, public disability insurance (DI) is one of the few relatively generous federal social insurance programs, even in an international comparison, see Besharov and Call (2022). As a result of rising recipiency rates and spending, researchers have analyzed the implications for labor supply, earnings, beneficiary health and well-being, multi-generational "welfare" cultures as well as household income, consumption and poverty (Dahl et al., 2014; Gelber et al., 2023; Autor et al., 2019). Using quasi-random case worker assignment, studies inside and outside the United States conclude that employment rates among marginally rejected applicants are 10 to 30 percentage points higher compared to marginally accepted applicants (Bound, 1989; Chen and Van der Klaauw, 2008; von Wachter et al., 2011; Maestas et al., 2013; French and Song, 2014; Kostøl and Mogstad, 2014). Further, the generosity of the public DI system and the stringency of the health screening process are major determinants of the inflow of cases (Autor and Duggan, 2003; De Jong et al., 2011). Finally, it is a stylized fact that receiving DI benefits is usually an absorbing state, which is why reform debates often surround the question of how to prevent DI take-up in the first place. For example, Burkhauser and Daly (2012) propose experience-rated premiums to incentivize worker accommodations after health shocks. Other U.S. policy reform proposals also target a better accommodation and suggest partial disability benefits to achieve that goal (Autor and Duggan, 2010; Maestas, 2019).

In this paper, we study the short and long-run consequences of a fundamental reform to the German public DI system. It became effective in 2001 and cut an entire category of insurance benefits for younger cohorts. We first study how the reform affected public DI inflows in the short and long-run. Then, we study how it affected the private individual market for disability insurance in Germany. Studying interaction effects between public social insurance systems and their private analogues are a key area of research in economics. Interaction effects allow researchers to infer how individuals value insurance (Cabral and Cullen, 2019) or to analyze the potential for private markets to substitute for reduced government provision of benefits.

In the U.S., rich research analyzes selection and crowd-out effects between Medicaid (means-tested, state-level health insurance) and private health insurance markets (Cutler and Gruber, 1996; Card and Shore-Sheppard, 2004; Clemens, 2015). The relevance of Medicaid—the only public insurance that covers long-term care expenses—for private long-term care insurance markets has also attracted economists' attention. The private market is fairly small despite a high long-term care risk in old age (Sloan and Norton, 1997; Brown and Finkelstein, 2008; Braun et al., 2019). In very recent work about choice decisions between the original and the privatized strand of Medicare, Cabral et al. (2023) find that individual factors explain more than half of the geographic variation in private take-up. Ours is one of the first papers to study interaction effects between a federal entitlement disability insurance program and an individual private long-term insurance market for work disability risk. While some published papers have described characteristics of the private German DI market (Soika, 2018; McVicar et al., 2022)), the economics literature on private markets for disability insurance is very small, with Autor et al. (2014) being a notable exception.[1]

We study a DI reform in the context of the German welfare state, which is known to be generous but whose public DI system became significantly less generous over time, see Section 2.2 as well as McVicar et al. (2022) for an comparative overview of all reforms since the 1970s. The 2001 reform studied in this paper substantially reduced the generosity of the public DI system for those born after 1960. We call those cohorts the "German notch cohorts." Effective 2001, the notch cohorts lost access to public "Occupational Disability Insurance (ODI)." ODI insures the lifecycle risk that a health shock prevents someone from working *in their previous occupation* (or a comparable occupation in terms of income and social standing). Thus, ODI can be thought of as a supplemental insurance that tops up the basic work "Work Disability Insurance (WDI)" which insures the risk to become work disabled in *in any job*. Besides Burkhauser et al. (2016), Börsch-Supan et al. (2022) and McVicar et al. (2022), we are not aware of published economic research that has assessed this reform.[2]

[1]In his unpublished job market paper, Seitz (2021)) estimates a dynamic life-cycle model to conclude that, with a coexisting private market, the welfare-maximizing public DI program is less generous than in a world without private markets. This paper also contributes to research studying interdependencies of coexisting social insurance systems and spillover effects between those (Borghans et al., 2014; Lalive et al., 2015; Koch, 2015; Leung and O'Leary, 2020; Ahammer, Alexander and Grübl, Dominik and Winter-Ebmer, Rudolf, 2020).

[2]In a concurrent working paper, Seibold et al. (2022) use data from a single DI insurer to study the same reform. In line with our findings using representative data, post-reform for affected cohorts, they find significant but modest

Using 1995-2018 administrative data from the German Statutory Pension Insurance (*Deutsche Rentenversicherung, DRV*), we show that the reform significantly reduced the inflow of new DI recipients, on average by 20% among males and 10% among females from 2001 to 2018. In 2011, when the first notch cohort turned 60, the long-term decrease stabilized at roughly 35%. We validate this substantial reduction in the public DI recipiency rate using representative household panel data from the German Socio-Economic Panel Study (SOEP).

In the next part of the paper, we study interaction effects with the private individual market for disability insurance in Germany. The German private individual disability insurance market is one of the biggest worldwide. It is almost exclusively an individual long-term market for ODI policies that tops up public benefits. Moreover, it provides experience-rated policies without guaranteed issue and resembles the private long-term care insurance market in the US (Brown and Finkelstein, 2008). We first characterize this market descriptively and discuss its regulation. In a Regression Discontinuity (RD) design with representative data and birth year as the running variable, next, we do not find much evidence that the notch cohorts purchased private ODI policies at differentially larger rates, compared to control cohorts. Then, we carve out three stylized empirical facts about this understudied market that we subsequently explain using the general equilibrium model: (1) generally low take-up rates, in particular given the high lifecycle risk of work disability, and relatively small substitution effects with the public DI market. (2) Strong income and health gradients in take-up, and (3) an—inversely related—higher risk of work disability among low-income and health risk groups.

After that, building on Braun et al. (2019), we use a customized version of their general equilibrium model and inputs from various data sources to analyze the role of three key factors in producing these stylized market equilibrium outcomes. The three factors are: (i) the German means-tested basic income safety net program, (ii) administrative costs, and (iii) private information. Finally, we provide policy simulations. We are asking: "What factors could policymakers target to increase take-up rates, given the current regulation?" In this market, targeting the high level of administrative costs appears to be the most effective and feasible policy option.

increases in private ODI sales by this insurer and a low "observed willingness-to-pay of many individuals." They argue that distributional concerns and potentially biased perceptions by workers would imply that the reform was welfare-decreasing.

2 The German Disability Insurance System

2.1 Social Insurance in Germany

In an international comparison, Germany has a generous social safety net consisting of public Unemployment Insurance (UI), Workers' Compensation (WC), Health Insurance (HI) and Long-Term Care (LTC) insurance (Schmieder et al., 2016; Bauernschuster et al., 2020; Fischer and Korfhage, 2023). Among employees, eligibility for sick and medical leave is universal (Ziebarth and Karlsson, 2010, 2014; Ziebarth, 2013).

Moreover, Germany runs a Statutory Pension Insurance (SPI) program (Eibich, 2015; Geyer, 2021) that includes the public DI program (more details below), and also a universal means-tested basic income cash transfer program. The means-tested social safety net program provides a guaranteed minimum income floor of about $1,000 per month for a single individual. For those who are able to work, it is called Unemployment Insurance II (*Arbeitslosengeld II*).[3] For those who are unable to work, it is called Social Assistance Benefits (*Hilfe zum Lebensunterhalt*) and has no job search requirements; recipients are not part of the labor force (§§27-40 SGB XII). In the final part of the paper, we will analyze the role of this basic income safety net program for low private ODI take-up among low-income individuals.

These social insurance programs are funded through a mix of contribution rates for UI, WC, HI, LTC and SPI as well as employer mandates for paid sick leave and general taxes for the social minimum income floor. See Eichhorst et al. (2008); Ziebarth (2018); McVicar et al. (2022) for more detailed overviews.

2.2 History of Public Disability Insurance in Germany: 1970 to 2001

Germany's public Disability Insurance (DI) program is part of SPI. It provides benefits for both partially and totally disabled workers, who have paid contributions during their work lives. Employers and employees are each subject to a payroll tax (since 2018: 9.3%) of their monthly gross wage up to the social insurance contribution ceiling of €7,050 per month (in 2022).[4]

Appendix Figure A1 shows the development of Germany's public DI caseload from 1970 to 2018 along with select reforms. Note that the figure shows the *stock* of all recipients. As such,

[3] A structural reform in 2004 streamlined and re-redesigned those programs. It introduced the *Arbeitslosengeld II* program, decoupled means-tested benefits from previous income, and cut the maximal duration of standard UI benefits, see *Social Code Book II*. For more information about the reforms see, Eichhorst et al. (2008); Konle-Seidl (2012); Dustmann et al. (2014). The reforms did not differentially affect the notch cohorts but cut the generosity of these alternative social insurance routes generally.

[4] The contribution ceiling is lower in East Germany at €6,750.

even large declines in the inflow of new recipients only gradually translate into overall DI rate declines.

In the early 1970s, compared to other OECD countries, Germany had high disability recipiency rates (Burkhauser et al., 2016). In 1972, a major welfare expansion introduced new early retirement benefits without actuarial deductions. DI enrollment rates kept on rising, peaking at 5.8% of the workforce in 1984. In 1982, the newly elected center-right government restricted eligibility to workers who had paid SPI contributions over the past three out of five years. As many housewives (and househusbands) did not meet these criteria, the strong decline in DI recipiency rates between 1984 and 1990 has been primarily linked to restricting access for women without much formal labor market attachment. Börsch-Supan and Jürges (2012) provides a more detailed discussion. In 1996, reforms introduced caps on the allowed labor market earnings of recipients. Subsequently, a strong reduction in the inflow of new recipients contributed to the decline in the recipiency rate over the rest of the decade (Figure A1).

2.3 The Fundamental Public Disability Insurance Reform of 2001

Until 2001, the German DI system included two schemes: (a) the basic work disability insurance (WDI) and (b) the supplemental occupational disability insurance (ODI).

WDI eligibility implied an earnings capacity test. The applicant must not be able to earn more than 640 DM (about €320 in 2001 or $480 today). ODI eligibility implied a reduction in work capacity "to less than half of that of a physically, intellectually, and mentally healthy person with similar training, knowledge and abilities" (§43 of Social Code Book VI; Viebrok (2018)). Importantly, the applicant must be "unable to work in the occupation in which one was trained—effectively in the last job or a comparable job in terms of the skills it required, the wages it paid and its prestige." While WDI provided full disability benefits, ODI only provided two thirds of the full benefits. The idea of ODI as a supplemental DI benefit was to compensate for *partial* loss of work capacity, hence the lower benefits. It was never intended as a benefit to solely provide the means for a living but as a social insurance compensation when a health shock required that someone downgraded from a full to a part-time job or a higher to a lower paid job. However, if reasonable part-time work was unavailable due to the local labor market situation, full benefits could be granted (Viebrok, 2018). [5] Appendix B discusses and simulates ODI benefit

[5] In other words, the local labor market situation mattered, especially when applicants could not be referred to another "reasonable" job, following a hierarchical scheme of four categories where workers could be referred to a job "one degree below" their actual category. In practice, pension case workers would ask the UI office if part-time jobs were available.

calculations in the context of the SPI system. Pre-reform, the average simulated replacement rate was 18% of average gross earnings. In practice, in 2000, the average public ODI benefits were €466 (or about $500) per month.

Since January 1, 2001, for cohorts born after January 1, 1961—our treated "notch cohorts"—public ODI was abolished. By contrast, those who were already 40 at the beginning of 2001 were grandfathered in and still had access to public ODI. We use the grandfathered cohorts as our control group. The notch cohorts are our treatment group.[6]

Health Assessment. The reform implemented more changes which, however, did not affect cohorts differentially. The new DI system, instead of using an earnings capacity threshold, measures work capacity by working hour capacity. If applicants' work capacity lies between 3 and 6 hours per day, then partial WDI is granted (50% of full benefits).[7] If applicants' work capacity is less than 3 hours per day, full WDI is granted (Deutsche Rentenversicherung, 2020). Moreover, since the 2001 reform, benefits are granted for an initial period of three years and have to be re-certified. After nine years, the benefit becomes permanent. If the work capacity is not expected to improve, a permanent pension can be granted earlier, which applies to half of all new cases.[8]

Eligibility Work Requirements. Importantly, however, the main work requirements to establish eligibility did not change in the course of the 2001 reform. Applicants must have paid SPI contributions in the last three out of five years. There is also a general waiting period of five years.

Application procedure. Details of the application and health assessment are specified in German Social Law and Deutsche Rentenversicherung (2018). Applicants apply at a SPI field office. They have to to submit all substantiating documentation such as medical diagnoses as well as medical tests and reports. Then, an independent and third-party medical doctor who is certified to

[6] In the course of the reform, an entirely new *Social Code Book IX* was passed. It regulates Rehabilitation and Participation in Social Life (*Rehabilitation und Teilhabe Behinderter Menschen*) for disabled and handicapped people in Germany. Before 2001, most of these regulations were included in the *Schwerbehindertengesetz*.

[7] As with ODI, partial WDI intends to compensate for a partial work capacity loss. However, if part-time jobs are unavailable, it can be converted to full WDI. This is usually assumed to be the case if the WDI recipient cannot find part-time work within a year. The share of partial WDI benefits converted to full WDI benefits has lied between 6 and 16% between 2001 and 2021 (Deutsche Rentenversicherung, 2022). Earlier data are not available. For grandfathered cohorts, the only criterion for ODI is whether applicants could work 6 hours per day in the previous occupation.

[8] Another reform in 2004 mandated employers to provide *Workplace Reintegration Management* ("Betriebliches Eingliederungsmanagement", §84 SGB IX). The idea is to overcome temporary disability and to prevent future deteriorations in work capacity. However, this reform is beyond the focus of this paper and affected all birth cohorts equally. It likely had a gradual impact on the decreasing stock of DI recipients as seen in Figure A1.

carry out medical assessments for the SPI reviews the case. Sometimes these reviewers are state-employed physicians (*Amtsärzte*), and sometimes they are regular specialists practicing in the county of residence of the applicant. In any case, the medical evaluators must not have any pre-existing relationship with the applicant. It is worthwhile to note that 44% of all applications are rejected; this share has remained stable since 2000.

Benefit Calculation. When approved, recipients receive benefits as a type of "early retirement pension" with actuarial reductions. Benefits are a function of recipients' earnings histories and not adjusted for family composition, income or assets. They are calculated as old-age pensions, assuming that recipients' would have earned their pre-DI labor market income until age 60. Further, actuarial deductions of 3.6% per annum are applied but capped at 10.3% for everyone receiving benefits before age 57. [9] Before 2001, ODI benefits were two thirds of full benefits. After 2001, for the grandfathered cohorts, this level was reduced to 50%. Obviously, this reduced benefit level decreased the relative attractiveness of applying for ODI. Thus, our reduce-form estimates on how abolishing public ODI affects inflows represents a lower bound estimate. We come back to this aspect in the last part of the paper and in Appendix B.

In 2020, the average benefit among new recipients was €882 (about $1,000) and 87% of those were full, not partial, benefits (Deutsche Rentenversicherung, 2021). Appendix B discusses the details of the benefit calculations and simulates replacement rates.

2.4 Private Disability Insurance in Germany

The German private disability insurance market is overwhelmingly an individual market, not a group market like in the United States (Autor et al., 2014). Similar to the long-term health insurance market in Germany (Atal et al., 2019, 2020), the private individual ODI market is individually underwritten. Guaranteed issue does not exist. Private disability insurance follows private insurance law (*Versicherungsvertragsgesetz*). It is based on a private contract between the insurer and the insured which specifies conditions for the insured risk. Premiums depend on age, medical diagnoses, health behavior, income, and occupation. As a result, premiums can easily be several hundred dollars for high-risk occupations and, often, applicants are denied coverage.

Rating agency data from competing private insurers covering almost five million ODI policies

[9] Several studies document an increased risk of poverty among people on WDI (Krause et al., 2013; Märtin et al., 2012, 2014; Geyer, 2021). As a consequence, policymakers increased WDI benefits again by increasing the "reference age" to 62 in July 2014, and to 65 years and 8 months in 2019. Now it equals the statutory retirement age and will further increase to 67 years by 2031.

shows that, in 2019, 23% of new applications were either rejected (8%), included pre-existing condition clauses (11%), or included risk-premia (4%) due to pre-existing conditions. Note that these are *conditional* on applying for a policy. In reality, brokers and online calculators easily tell potential applicants in advance whether an application has some chance of success or not. In 2014, a highly respected consumer magazine reported that 235K applications per year would be rejected by the industry, and revealed that 81% of those who were offered a policy were offered a less generous coverage than desired (Ökotest, 2014).

The average age when signing a policy is 32, but the age distribution is left-skewed with 64% of new policyholders being below 31. The average age when becoming work disabled is 46, and the average contract runs until age 64. In 2019, the four main reasons that triggered an approved occupational disability in the private market were: mental diseases (32%), musculoskeletal diseases (20%), cancer (18%), and accidents (8%).

In our representative data, focusing on those between the age of 20 and 59, about a third of all households with an employee or self-employed person as household head were covered by private disability insurance, which is almost always ODI coverage (Statista, 2014). In 2015, according to the German Association of Insurers (GDV), the average pension from a private ODI was at €629 per month (about $750 in 2022).

The ten biggest insurers hold more than 60% of the total market share and offer similar but not identical policies. The market is characterized by freedom of contract between insurers and applicants. Many online calculators yield advice on a wide range of policy elements that can be individually customized leading to several hundreds of contracts. In an audit study, Ökotest (2014) found very large differences in premiums by occupancy and health. Together with healthy profit margins, these facts suggest monopolistic market structures.

In contrast to the U.S. market where private group DI usually includes "offset clauses" that may reduce public Social Security Disability Insurance benefits dollar for dollar (Burkhauser and Daly, 2012), in Germany, private and public DI benefits do not crowd each other out. In fact, they are independent and private ODI benefits top up public benefits. Further, the private insurance industry relies on their own medical examiners and there is no coordination between SPI and private insurers (BBP, 2020).

In the last part of the paper, we will further characterize the private ODI market, carve out several stylized empirical facts, and then use a general equilibrium model to study the role of (i) the means-tested basic income program described above, (ii) administrative costs, as well (iii) private information for equilibrium market outcomes such as coverage denials and market selection.

3 Impact of the 2001 Reform on Public DI Inflows and Case Loads

In a first step, we provide evidence on the first-stage effects of the fundamental public DI reform. That is, we show how it affected the inflow of new recipients and the case load. To do so, we use two types of datasets and two reduced-form identification approaches: (1) an administrative dataset on the inflow of cases by birth cohort and year in a difference-in-differences (DD) framework and (2) representative household panel data from the German Socio-Economic Panel Study (SOEP) in a regression-discontinuity (RD) design. We use (2) to validate the findings in (1) using the universe of the underlying cohort populations, not just select inflows.

3.1 Impact on Public DI Inflows Using Administrative Data

First, we use administrative data from the SPI (*Deutsche Rentenversicherung, DRV*) to estimate the impact of the 2001 reform on the inflow of public disability insurance (DI) recipients. The data are available by year, region, gender and birth year.

DD Method. We normalize the number of inflows by cohort for each year using population data from the Federal Statistical Office.[10] Further, we focus on cohorts born between 1954 and 1966 and the ages of 29 and 59 in a given calendar year. Then, using data from 1995 to 2018, we compare our treatment group—the notch cohorts who were ineligible for public ODI from January 2001—to our control group—the grandfathered cohorts who were born before 1961. In robustness checks, we also vary the bandwidths. We estimate the following Difference-in-Differences (DD) model:

$$y_{ct} = \alpha + \beta D_c \times T_t + \delta_t + \rho_c + \epsilon_{ct} \qquad (1)$$

where y_{ct} denotes the share of new public DI recipients of cohort c in year t; D_c is a dummy that identifies notch cohorts; T_t is a post-reform indicator that turns on after 2000; δ_t are year fixed effects; and ρ_c are cohort fixed effects. ϵ_c denotes the error term, which we cluster at the cohort level.

The main identification assumption implies that, absent the reform, the inflow of new public DI recipients of the notch cohorts would have developed in the same manner as those of the

[10] We use unconditional shares, i.e., we do not subtract the number of people currently receiving DI.

grandfathered control cohorts. Note that our setting is not prone to possible biases due to effect heterogeneity as in staggered DD settings (Goodman-Bacon, 2021).

Results. To illustrate the main findings, Figure 1 plots an event study using equation (1) but replaces T_t with a series of year dummies, where 2000 serves as the baseline year.

[Insert Figure 1 about here]

As seen, whereas the five pre-treatment years show no trending and relative inflow differences between treated and control cohorts are not significantly different from zero, we observe a sharp decline in inflows beginning in the first post-reform year 2001. This decline further accelerates in subsequent years, up to point estimates exceeding -0.2 percentage points, or about 35% relative to the pre-reform mean.

By 2011, one decade after the reform implementation, the inflow differential between the two groups had flattened out. From then on, it remained highly significant at -0.2 percentage points. This represents the long-run effect of the reform.[11] Recall that ODI benefits also decreased for the grandfathered cohorts, see Appendix B for details. To the extend that these reduced benefits significantly affected the likelihood to apply for DI, our estimates here represent lower bound estimates.

Figure A2 (Appendix) shows the same event studies separately by gender. Again, we observe reassuringly stable pre-reform trends, followed by substantial inflow reductions among the notch cohorts. However, not surprisingly, the reform-induced decrease in inflows is substantially larger for males. The reason is that their eligibility rates are higher due to a stronger labor market attachment. Specifically, men are more likely to fulfill the eligibility requirement and have paid pension contributions during the last three out of five years. Moreover, men are more likely to work in physically demanding occupations and industry jobs that generally carry a higher work disability risk.

Table A1 (Appendix) shows the DD regression model equivalents, where Panel A shows results for the full sample, Panel B shows results for men, and Panel C shows results for women. Each column in each panel stands for one separate DD model like in equation (1).

The findings in Table A1 are in line with the event study estimates. First, the estimates are robust to the inclusion of cohort and year fixed effects as well as controls for East Germany. The average decline in inflows for males translates into a 20% decrease, relative to the mean of the

[11] For this number, we use as pre-reform mean the mean entry rate of untreated cohorts which was 0.58% for cohorts born between 1954 and 1960.

control group. The decline for women is only half as large at 10%. However, when zooming-in and restricting the bandwidths of the cohorts considered, that is, cohorts born in 1959 to 1962, the effect sizes decrease to -12.5% for males and -7.9% for females—on average over *all* post-reform years 2001-2019. Note that the long-term effect from 2011 onwards is about twice as large (see Figure A2). Also note that, reassuringly, these reform effects mirror the pre-2001 share of ODI pensions among all new recipients. Viebrok (2018) reports relatively stable shares of between 12 and 18% for men and about 8% for women in the 1980s and 1990s among new recipients.[12]

3.2 Impact on Public DI Case Load Using SOEP Survey Data

Data. In a second step, we validate our first-stage findings above using representative household data from the German Socio-Economic Panel Study and an alternative identification approach. The SOEP allows us to observe representative samples of the entire cohort, not just inflows as with the administrative data, see Goebel et al. (2019) for more details on the SOEP. We relegate most details to the appendix and focus on the main approach and findings.

Sample Selection. We focus on the years 1995 to 2016 and respondents between the age of 25 and 59 (as we can then unambiguously identify whether they receive a public disability pension). In addition, we focus on birth cohorts from 1950 to 1970. Table A2 shows the summary statistic, where we list our main outcome variables in the upper panel and the covariates in the lower panel.

RD Method. As we are now using the underlying population of interest, we are able to study the impact of the 2001 reform on public DI case loads using a Regression Discontinuity (RD) design. The discontinuity is the birth year 1961. It determines whether respondents belong to the notch or the control cohorts. A standard linear parametric RD model is:

$$y_{it} = \alpha + \beta D_i + \psi(1 - D_i) f(z_i - c) + \gamma D_i f(z_i - c) T_t + X'_{it}\tau + \delta_t + \rho_s + \epsilon_{it} \tag{2}$$

where y_{it} indicates whether the respondent receives public DI benefits. D_i is one if the respondent belongs to the notch cohorts. The cohort measure z_i enters in difference to the reform cutoff c, 1961. Including linear trends and polynomials in the running variable $f(z_i - c) = z_i - c$ allows for different slopes before and after the cutoff.

All regressions include year (δ_t) and state (ρ_s) fixed effects. X'_{it} represents a rich set of socio-demographic, educational and job-related control variables as listed in Table A2. For example,

[12] However, recall that many ODI pensions were converted to full WDI pensions if recipients could not be referred to a “reasonable” job, see Section 2.

45 is the average age, 52% are women, and 71% are married. About 20% finished the highest educational track in Germany and 21% are part-time employed; 42% are white-collar employees.

We follow the recent literature on the topic and do not cluster standard errors ϵ_{it} (Cunningham, 2021). Further, we follow the literature and estimate nonparametric local polynomial regressions with univariate weights and cubic terms as our baseline model (Calonico et al., 2014). We present robust and bias-corrected estimates (Calonico et al., 2018), vary the bandwidth, use data-driven bandwidth selection (Calonico et al., 2020), and covariates (Calonico et al., 2019) in the Appendix.[13] Moreover, our estimates are robust to implementing methods for discrete running variables following Kolesár and Rothe (2018).

Despite all econometric sensitivity checks, the main RD identification assumption implies that no other factor would have affected public DI caseload trends discontinuously at the birth year level. We are not aware of another reform or factor that could invalidate this assumption; the Appendix provides further evidence that other covariates trend smoothly at the cut-off c.

Outcome. The SOEP Group provides a time-consistent longitudinal binary variable that indicates whether individuals receive an old-age pension due to work disability. We call this variable *Public DI I*. Moreover, the SOEP Group provides a second generated variable indicating the annual income stream from old age, disability or civil servant pensions, which we use to create a second binary indicator, *Public DI II*.[14] According to Table A2 and *Public DI I*, 3.3% of the German working age population have been on DI between 1995 and 2016—this share matches the share from official data in Figure A1 very well.

Results. Figure 2 plots public DI recipiency rates by birth cohorts. It displays unconditional scatters by year of birth, overlaid with polynomial quadratic smoothing plots. The visual evidence from the representative SOEP corroborates the findings from the administrative data: we see a clear discontinuous decrease in the probability of receiving a public DI pension for the notch cohorts in post-reform years. FigureA3 shows no such discontinuity for the pre-reform years in the left column. Moreover, using either *Public DI I* (first row) or *Public DI II* (second row) yields robust findings.[15]

[Insert Figure 2 about here]

Table A4 shows the RD results using local polynomial RD methods for the post-reform period

[13] We also implement procedures for optimal local polynomial order selection following Pei et al. (2022).

[14] Here we use only respondents with a positive pension amount who do not receive a civil servant, a veteran's, a miners' or a farmers' pension.

[15] Note that the DI level is higher for post-2001 years as our respondents are older compared to 1995 to 2000. The decreasing slopes imply decreasing DI rates by birth cohort.

from 2001 to 2016. The column headers indicate the outcome measure; the lower panel adds socio-demographic and educational covariates as indicated. The models in columns (3) and (4) use *Public DI I* but restrict the sample to non-married respondents and single households, respectively. The table shows the results from 24 different models; for each column and panel, we present results from conventional, bias-corrected and robust RD models, see Calonico et al. (2014, 2017, 2019) for details.

As seen, we find statistically significant results for 22 out of 24 models; all 24 models produce consistently negative point estimates, in line with Figures 2 and A3. Our preferred bias-corrected and robust estimates of the first column are -1.6 percentage points (upper panel) and -1.5 percentage points (lower panel). Relative to the mean recipiency rate of the non-treated cohorts, 6.7%, the latter estimates translate into a decrease of 22%. The size of the decrease for households with one member is very similar, whereas the decrease for non-married people is even larger. Overall, the findings confirm and validate the results from administrative data that just focus on inflows.

The Appendix shows the results from various robustness checks. Figure A4 varies the bandwidth and also uses data-driven bandwidth selection methods (Calonico et al., 2020); Figure A5 shows that covariates such as age, children in the household, white collar, or Self Assessed Health (SAH) trend smoothly at the cutoff 1961. Figure A6 carries out a McCrary (2008) density plot of the running variable, and Figure A7 varies the polynomial (Pei et al., 2022), the weights, runs donut RD models, and adds a full set of covariates (Calonico et al., 2019).

Pre- and Post-Reform Consequences of a Health Shock. Did relevant outcomes linked to health shocks shift for the treated, the younger generation, in post-reform years? In other words: For the younger generation without access to public ODI, how does a health shock materialize, given other social insurance strands and intra-household risk sharing?

To investigate this question, Table A6 (Appendix) uses SOEP data from 2001 to 2016 and runs standard individual fixed effect OLS models. Each column is one model that includes as (lagged) regressors the same indicator for severe health limitations as above[16], a dummy for whether respondents belong to the treatment group (born after 1960) as well as the interaction between the two. The dependent variables are whether, in the subsequent year, (1) the respondent is on public DI, (2) the respondent is not employed, (3) the respondent's total market and non-market income as well as (4) her subjective well-being.

[16]However, here we use an annual binary indicator for several health limitations whereas we use the cumulative lifecycle risk in Figure 5

As seen in Table A6, the onset of a severe health limitation more than doubles the likelihood to be on public DI in the next year (column (1)) and, by the same share of 9ppt, increases non-employment. Further, total annual income decreases significantly by €4.2K (-14%, Table A2) as does subjective well-being (-0.18 points on a 0-10 Likert scale). Moreover, while the interaction term between the health shock and the treatment dummy yields a point estimate in line with the effects in Figure 2 and Table A4, it is imprecisely estimated. Similarly, the interaction effects suggest (imprecise) increases in non-employment by about 4ppt, and small and insignificant effects for income and well-being changes. Overall, there is not much evidence that the treated cohorts did substantially and significantly worse in terms of income and well-being as a result of health shocks compared to the control cohorts.

4 The 2001 Public DI Reform and the Private DI Market

In the second part of the paper, we study interaction effects between the public and private DI market. Moreover, we provide general insights into the German individual private DI market, one of the biggest in the world. Specifically, we carve out several empirical stylized facts in the context of its market regulation (see also Section 2). Then, building on Braun et al. (2019), we use a general equilibrium model to explain these stylized empirical pattern. Specifically, the model has the power to leverage three main elements—private information, administrative costs, and the German means-tested basic income program—to replicate various empirical market pattern: (1) Given a high lifecycle risk of work disability, private ODI take-up rates are relatively low and unresponsive to losing public ODI. (2) Private ODI take-up rates increase strongly in good health and income, and (3) are inversely linked to the lifecycle work disability risk.

4.1 Impact of the 2001 Reform on the Private ODI Market

In a first step, we investigate whether the notch cohorts purchased private ODI policies at higher rates, relative to the non-notch cohorts, to compensate for the loss of public ODI coverage. It is a straightforward hypothesis that the reform may have crowded-in demand for private ODI. A rich economics literature has studied the reverse effect, *crowd-out* of private health insurance through public health insurance expansions (Cutler and Gruber, 1996; Clemens, 2015).

This is one of the first studies to estimate the impact of *reductions* in public social insurance generosity on the market for private insurance. It is also one of the very first papers to study interaction effects between public and private DI markets, see Cabral and Cullen (2019) for a rare exception of published work.

Data. For this exercise, we rely on representative survey data from the SAVE survey (Saving for Old Age in Germany, *Sparen und AltersVorsorgE in Deutschland*). Coppola and Lamla (2013) provide a detailed overview of the dataset. The SAVE data include a very rich set of questions about preferences, savings, retirement, health as well as standard socio-demographics. Some of these measures are typically unobserved by researchers and insurers. This unique survey helps us to (a) mimic the risk classification of private ODI insurers and to (b) assess private information that drives insurance market selection in the spirit of Akerlof (1970) and Hendren (2017).

Sample Selection. We use all SAVE waves from 2001 to 2010, which were conducted annually (except for 2002 and 2004). We again focus on employees below the age of 60.[17] Table A3 shows the summary statistics of our main sample. 32% of all households are ODI policyholders, the average age is 41 and 41% earned the highest schooling degree in Germany after 13 school years. To identify the notch cohorts, we directly observe the birth year as a separate variable.

Figure 3 illustrates the main result for the full sample; Figure A8 (Appendix) shows robustness checks for alternative samples; clockwise, starting from the upper left: the full sample as in Figure 3, respondents who are eligible for public DI, childless households, and one-person households. The x-axis displays the birth year, and the y-axes display the outcome variable, *Private ODI*. We again plot unconditional scatters by birth year, overlaid with linear plots for each side of the cut-off.[18]

[Insert Figure 3 about here]

The figures show the following: First, the demand slope is clearly and strongly increasing in the birth cohorts. In other words, younger people are much more likely to be covered by a private ODI policy in Germany. This observation is not surprising. The reason is that, after a strong expansion of the welfare state in the decades after WWII (especially in the 1970s), German policymakers started to implement a series of structural reforms of the statutory pension and DI system in the 1980s and, to a great extent, the 1990s, see Figure A1. The structural reforms in the second half of the 1990s and early 2000s were accompanied with especially strong messaging, education (also in schools) by consumer advocates, and lobbying that private insurance policies for old age protection would be crucial for young people. In addition to shifts in the public

[17]We ignore civil servants who were not affected by the DI reform.
[18]Linear slopes fit the data better than quadratic ones; however, we vary polynomials in robustness checks.

perception of the importance of private insurance to cover future life shocks, younger cohorts are much less likely to be rejected by private ODI insurers and are offered lower premiums as they are healthier and have fewer pre-existing conditions, see Section 2.

Second, none of the figures shows an obvious discontinuous jump in the likelihood to have private ODI insurance for the notch cohorts. While single insurers may certainly have targeted subgroups that were affected by the 2001 reform (Seibold et al., 2022), representative data do not yield much evidence for a systematic and substantial crowding-in or substitution effect.

Table A5 shows the equivalent local polynomial RD results for Figure 3, following the same table setup as above. As seen, three of the four sample specifications with the associated 18 models show consistently non-significant point estimates. For example, for the full sample in column (1), we obtain bias-corrected and robust RD estimates of size 0.05. Overall, in Table A5, 19 out of the 24 estimates carry negative signs, not the hypothesized positive ones. The model that focuses on those eligible for public pensions (and thus DI) produces negative and statistically significant point estimates when including the full set of socio-demographic and labor market controls.

Robustness checks vary the bandwidth (Figure A9, Calonico et al. (2020), study discontinuities in covariates (Figure A10), plot the density of the running variable (Figure A11, McCrary (2008), and alter polynomials (Figure A12). These use our preferred model in column (1) with exogenous controls (age, gender, year and state fixed effects) and do not yield any evidence for positive and statistically significant effects. Further, correcting for the discrete running variable (Kolesár and Rothe, 2018) does not alter the findings (detailed results available upon request). However, the robustness checks also illustrate that most point estimates carry relatively large standard errors. Nevertheless, we can exclude with 95% statistical certainty that the notch cohorts took up private ODI insurance at a rate higher than 12 percentage points (ppt), relative to the baseline of 33% (Table A5) as a result of the reform.

4.2 Some Stylized Facts on the German Individual Private ODI Market

While there may be higher differential take-up of private ODI policies among subsamples or among single insurers, apparently, there is not much evidence for systematic, strong and significant increases in the general population. Even when considering the upper 95% bounds of the statistical confidence interval of our preferred specification in column (1) of Table A4, the increase in take-up was at most 12 percentage points off a baseline of 32%, leaving the majority of

German employees uninsured. This begs the question "why is that?"

One possible interpretation could be a lack of demand, which may imply that people do not value ODI coverage highly. Under certain conditions, this would imply that the reform was welfare-improving. However, it should be kept in mind that the observed coverage outcomes are *equilibrium outcomes*. They are the result of an interplay between demand, supply and market regulation.

Thus, this section employs a general equilibrium framework based on Braun et al. (2019) to better understand and trace out underlying driving forces for the low post-reform take-up rates. We will begin by presenting several stylized facts about employee health, the lifecycle risk of becoming work disabled, as well as take-up in the private German market for ODI policies. In this part, we will refer back to Section 2 and further elaborate on the regulation of this market. To reiterate: Unlike in the United States where the private market for disability policies is mostly a group market (Autor et al., 2014), the German market is almost exclusively an individual market without guaranteed issue, pre-existing condition clauses and risk rating. It resembles the U.S. private market for long-term care and life insurance. The regulation and features mirror the German private long-term health insurance market, see (Atal et al., 2020) for further details. To provide these general empirical patterns, we rely again on the representative SAVE and SOEP surveys.

Health Risk Score. Table A3 shows a detailed list of health questions contained in the representative SAVE survey. For example, SAVE does not just feature the standard self-assessed health (SAH) measure but also a 0-10 Likert scale health satisfaction measure along with questions on health concerns and whether respondents have serious health issues. Further, it includes a list of the most common medical conditions such as heart disease, stroke, cancer, high blood pressure, high cholesterol, or chronic lung disease for each respondent. Smoking status is also sampled. Finally, SAVE elicits the number of doctor visits and hospital nights in the previous year. All these information reflect what private disability insurers ask in their health assessment questionnaires before making decisions about add-on premiums, pre-existing condition clauses or outright denials.

[Insert Figure 4 about here]

We use these information in conjunction with a principal component analysis to summarize and aggregate all available objective and subjective health measures into a continuous health risk score (Jolliffe, 2002)). The distribution of this normalized health risk score ranges between 0 and 1 and is in Figure 4. It is reassuring to see a typical left-skewed health risk distribution with a long right rail (Karlsson et al., 2016, cf.).

Next, we circle back to the representative household panel SOEP. The SOEP has existed since 1984 and allows us to leverage and trace out variation in the *lifecycle* risk to become work disabled. As mentioned, we obtained market-level data form a rating agency on the universe of contracts from 64 competing private insurers covering almost five million ODI policies. The average age when people purchase policies is 32, but the age distribution is left-skewed with 64% of new policyholders being below 31. The average age when becoming work disabled is 46, and the average contract runs until age 64.

Consequently, we use the SOEP to mimic these lifeclye pattern. We focus on a sample of respondents whom we observe at least once working full-time between the ages of 25 and 35, when Germans typically enter the labor market and decide on signing policies. We cut the SOEP lifecycle sample such that we also observe the same individuals at least once between the age of 55 and 60. By doing so and following Burkhauser and Schroeder (2007) in generating work disability measures using the SOEP, we elicit a broad and representative measure of the lifecycle risk of work disability among German employees.

[Insert Figure 5 about here]

Lifecycle Risk for Severe Health Limitations. Figure 5a plots this lifecycle risk of having a severe health limitation against the quintiles of self-reported health satisfaction at the age of 25-35. We use this measure of health satisfaction as a proxy for health when entering the labor market. [19] It also stratifies the risk by the quintiles of household net income. We summarize Figure 5a as follows: first, the lifecycle risk of a severe health limiting shock is large. Second, it remains significant even for the healthiest employees. It is 49% for those 20% with the lowest health satisfaction, then drops to 26% for the next quintile and further drops to 8% for those who are most satisfied with their health. Third, it entails a clear income gradient. It is 31% for the lowest income quintile, 20% for the second lowest, and then drops to 10% for the richest quintile.

Figure 5b shows the same graph but first traces out socio-demographics, job and educational characteristics (but not income and health). As seen, the curves flatten substantially over the baseline health status but maintain a clear income gradient. Further, the lifecycle risk remains

[19] Unfortunately, the SOEP only includes health satisfaction and the standard SAH measure over the whole 33 years that we use. The quintiles are not exact quintiles as they are derived from the 0-10 Likert scale.

high, above 20% for most health and income groups. For example, the lowest income quintile carries a work life risk of 37%, the second lowest of 24%, and the highest of 16%. All these pattern are very consistent with Meyer and Mok (2019) who report similar statistics for the United States using the PSID.

Stylized Facts on ODI Take-Up. Figure 6 summarizes some key stylized facts of private ODI take-up in a compact manner. The figure shows take-up on the y-axis and the population quintiles of the health risk score (Figure 4) on the x-axis. The downward sloping lines are again stratified by quintiles of net household income.

[Insert Figure 6 about here]

We can summarize: First, we see that take-up strongly decreases from the second lowest to the highest health risk quintile where a higher quintile indicates *worse* health. This pattern is not surprising, given that insurers can deny coverage and premiums are risk-rated. As discussed in Section 2, even conditional on applying for coverage, 24% of all policies are either rejected, contain a pre-existing condition clause or have health risk add-ons.

Second, Figure 6 shows that, across the entire health risk distribution, the lowest income quintile has take-up rates substantially lower than all other income quintiles, between 25% for the healthiest and below 10% for the sickest health risk quintile. In other words: The poorest 20% of the population have take-up rates of only 10 to 25%. The second lowest income quintile has also substantially lower take-up rates than quintiles three to five (but higher rates than the lowest quintile). For all income groups, we observe clear health gradients in take-up, meaning that take-up always drops significantly with worse health status.

In conclusion: (i) the lifecycle risk for severe health limitations is high—even corrected for socio-demographics and job characteristics—and between 15% and 40% for different levels of the health risk and income distribution. Further, (ii) this lifecycle risk increases with worse health and lower income. Nevertheless, (iii) private ODI take-up rates remain low, even after the substantial reductions in public DI generosity, and are between 10% and 50% for different parts of the health and income distribution. However, paradoxically, (iv) take-up is inversely related to the lifecycle work disability risk as the sickest and poorest have the lowest take-up rates despite having the highest work disability risk.

Finally, as before, Figure 7 plots private ODI take-up rates on the y-axis and the five risk score quintiles on the x-axis. However, the two lines differentiate by whether SAVE respondents expect to stop working before age 60 which proxies for expected work disability. As seen, over the entire declining health distribution, those who expect work disability have substantially higher take-up rates. While this empirical pattern is no definite proof of an adversely selected market, we interpret it as suggestive evidence for it.

5 General Equilibrium Model to Explain Stylized Facts

This section employs a variant of the general equilibrium model by Braun et al. (2019), which is based on Rothschild and Stiglitz (1976), Stiglitz (1977) as well as Chade and Schlee (2020). While the seminal Rothschild and Stiglitz (1976) model focuses on the role of private information to study adverse selection, every individual is insurable in this model. In addition to private information, Braun et al. (2019) enrich a monopolistic market model by adding administrative costs following Stiglitz (1977) as well as Chade and Schlee (2020). Chade and Schlee (2020) show that the existence of administrative costs can explain the empirically observed and economically relevant coverage denials to bad risks in insurance markets. In the standard adverse selection models, only good risks can go uninsured—voluntarily. Braun et al. (2019) develop a variant of Rothschild and Stiglitz (1976), as in Stiglitz (1977), to study the role of three key factors in insurance take-up: private information, administrative costs as well as Medicaid (public insurance for low-income individuals in the United States).

We build on Braun et al. (2019) and the literature above, but customize and adjust the model to capture the institutional details of the German private ODI market. The model by Braun et al. (2019) explains features and empirical puzzles of the U.S. private long-term care insurance market. However, it is powerful and flexible enough for us to tweak it to explain empirical equilibrium outcomes in the German private ODI market.

Specifically, our model leverages three main driving forces to explain low ODI insurance take-up after the fundamental public DI reform of 2001: (i) The German means-tested basic income program, (ii) private information, and (iii) administrative costs. Despite its simplicity, the model is powerful enough to capture the main regulatory framework of the private ODI market in Germany that leads to high denial rates for certain health risks, income and occupational groups. Appendix C discusses optimal contracts and market equilibria, given (i) to (iii). Note that the model is a standard general equilibrium model; it is not a behavioral model that explains low take-up rates by, for example, biased perceptions of work disability risk.

5.1 Quantitative Model

5.1.1 Individuals

From the empirical facts, see Section 2, we know that individuals buy private ODI insurance at an average age of 32. The average age when health shocks lead to occupational disability is 46, and the average contract runs until age 64, shortly before individuals hit the statutory retirement age of 65. Accordingly, an individual's decision-making problem entails three time periods as illustrated in Figure 8.

[Insert Figure 8 about here]

Period 1: Labor Market Entry and Endowments. In the first period, individuals enter the labor market between age 25 and 30. At the time, they draw a health endowment h, an economic endowment—consisting of wage w_1—and an occupation o. Individuals decide on how much to consume (c_1) and how much to save (s).[20] Health, wages and occupation are jointly distributed with density $f(h, w_1, o)$.

$$c_1 = w_1 - s \tag{3}$$

Period 2: ODI Offers and Purchase Decisions. While insurers observe health, wages and occupation, those are solely noisy indicators of the true work disability risk, $\theta^i_{h,w,o}$, with $i = b, t$. This true risk is an individual's private information. With probability $\rho = b$, individual i is at the bottom, and with probability $1 - \rho = t$ she is at the top of this probability distribution. Thus, the population share of those who incur a health shock which leads to work disability is $\eta \equiv \rho\theta^b + (1-\rho)\theta^t$.

The insurer operates in a monopolistic market, following Stiglitz (1977) and Braun et al. (2019), and maximizes profits, subject to participation and incentive constraints, see below. The insurer observes policyholders' h, w, o and either denies coverage or offers a menu of ODI contracts $(\Pi(h, w, o), b)$, where $\Pi(.)$ is the insurance premium and b are contracted insurance benefits in case of occupational work disability. As discussed in detail in Braun et al. (2019) and Chade and Schlee (2020), under the existence of fixed and variable administrative costs, insurers may deny entire risk groups coverage as they become unprofitable.

[20] At this time, educational decisions—a major driver of occupation and lifecycle income is completed for the great majority of the population (Carneiro et al., 2011; Atal et al., 2020). When using SOEP and SAVE data, we condition the empirical moments for Period 1 on individuals who we observe working full-time between age 25 and 35.

Individuals may or may not purchase ODI coverage that tops-up the basic public insurance due to several reasons.[21] First, in contrast to the insurer, they know whether their true disability risk is high or low, but they don't know their actual risk with certainty. They weigh the risk of occupational disability—which results in income losses up to the social insurance consumption floor C—against paying a monthly premium Π to insure this risk and receiving $C + b$ when their health prevents them from working in their previous occupation.

Second, the *participation constraint* (see below) ensures that each type i prefers, if offered, the specifically customized ODI policy over no insurance.[22] And the *incentive compatibility constraint* ensures that each type i prefers the specifically customized policy over the policy customized for the other type.

Finally, individuals also consider uncertainty about their future income, τ, that is unrelated to work disability, but may reduce or increase household income. The German social safety net provides a consumption floor that is means-tested at the household level to all residents. Thus, absent uncertainty toward the future rank in the income distribution, agents at the lower end of the income distribution have little incentive to buy private ODI insurance. Individuals in the upper end of the income distribution, on the other hand, may be eligible for means-tested basic income program in case of a negative shock to their household income (which is unrelated to work disability). It decreases the incentive to buy private ODI for this group. Also note that private ODI represents *supplemental* insurance for the potential income wedge between income from the last (or trained) job, relative to the lower paying job that individuals can still work in after a health shock leading to occupational disability. Thus, private ODI is less valuable, the smaller the potential income wedge.

Period 3: Income and Health Shocks. Period 3 represents the main work life of individuals and stretches from age 35 to retirement. In Germany, the earliest possible age to receive a statutory early retirement pension is 62.

As mentioned, individuals are aware that future labor market incomes w_2 are uncertain with density $q(\tau)$, where $\tau \in [\underline{\tau}; \bar{\tau}]$. A potential income shock may lead to eligibility for means-tested basic income that provides a consumption floor C.

[21] The "means-tested basic income program" and "public (W)DI" set very similar consumption floors and the former tops-up the latter if the WDP falls below the basic income level of roughly \$1,000 per month (Bundesagentur für Arbeit, 2019). The average monthly WDP pension for fully work disabled new recipients was €849 or about \$900 in 2019 (Deutsche Rentenversicherung, 2021).

[22] In the model, insurers would deny coverage if policies become unprofitable at reasonable premiums; hence, "only reasonable" policies are offered in this environment without guaranteed issue.

Moreover, individuals may experience a health shock that leads to occupational work disability. Those who experience such a health shock (partially)] lose their income. They either obtain C or a low-wage job (0.5*w) that essentially equals C by assumption.[23] Thus, in case of occupational work disability, those without private ODI incur costs $w_2 - C$, and those with a policy incur costs $w_2 - C + b - \Pi$ resulting in consumption c_{ODI}.[24]

$$c_{ODI} = (1-\tau)w_2 - w_2 + (1+r)s - \Pi + b + \Psi \tag{4}$$

where r is the real interest rate on savings, s, and Ψ is a social insurance transfer, for example, the German means-tested basic income or reduced income (0.5*w) whichever is higher, where $\Psi = max[0, C - ((1-\tau)w_2 + (1+r)s - \Pi + b)]$

Individuals that buy a private ODI policy but do not occur occupational disability, pay premia Π but incur no costs, resulting in consumption with an ODI policy but not health shock:

$$c_0 = (1-\tau)w_2 + (1+r)s - \Pi \tag{5}$$

In reality, the private ODI market provides hundreds (if not thousands) of different plans. Experts suggest to insure an income level of 70% of the gross wage, whereas various online calculators provide information on the trade-offs between premium and coverage levels as a function of h, o and w_2 (Allianz, 2022).

Consequently, individuals solve the following maximization problem, where we omit subscripts for readability:

$$U(h, w, o) = \max_{c,s,C} u_1(c_1) + \beta[\rho u_2(h, w_2, o, \theta^b, \Pi, b) + (1-\rho)u_2(h, w_2, o, \theta^t, \Pi, b)] \tag{6}$$

where

[23]This implicitly assumes either both occupational *and* work disability or that the next available job in case of occupational disability does not pay much more than C. Both are reasonable assumptions. First, occupational disability often implies work disability, especially for lower-paying jobs. Second, Germany has the biggest low-wage sector in Germany and those who have to downgrade their job due to occupational disability often earn below or not much above the consumption floor C which equals roughly 100% of the Federal Poverty Line (FPL) in the U.S. As discussed, cf. footnote 21, average DI benefits have consistently been below C.

[24]The average German gross wage was €47,928 or about $50K in 2019 (Statistisches Bundesamt, 2022). Thus, the average German who occurs a health shock leading to work disability at age 46 experiences an annual income loss of about $38K without ODI insurance. With an ODI policy at recommended levels of 80% of the gross wage, they their loss is $50-12+0.7*50=$3K p.a. for the 16 years between 46 and 62 (premia not considered). Average premia are currently €923 or $1000 per year according to our data from a big rating agency covering five million ODI policies, see Section 2. Hence, individuals spend on average $14K in premiums between ages 32 to 46 in case of a health shock. However, it should be noted that the actual coverage lies almost always below 80% of the gross wage; in 2014, a high-quality consumer rating report revealed monthly premiums between $50 and $200 for insured monthly benefits of between $750 and $2000.

$$u_2(h, w_2, o, \theta^i, \Pi, b) = \int_{\underline{\tau}}^{\bar{\tau}} u(\tau w_2) + \alpha[\theta^i u(c_{ODI}) + (1-\theta^i)u(c_0)]q(\tau)d\tau$$

where, α and β are discount factors.

5.1.2 Insurers

Applicants for private ODI indicate h, w, and o on their application, whereas the true work disability risk $\theta^i_{h,w,o}$ remains private information. The insurer either denies coverage or offers a menu of contracts $\{\Pi(h,w,o), b\}$ to profitable applicants. The insurer maximizes profits Ξ as follows:

$$\Xi(h,w,o) = \max_{\Pi,b} \rho[\Pi^b - \theta^b[\lambda b^b + \gamma I(b^b > 0))] + (1-\rho)[\Pi^t - \theta^t(\lambda b^t + \gamma I(b^t > 0))] \quad (7)$$

where, following and Chade and Schlee (2020)), variable insurer costs are λ and fixed insurer costs are γ. An example of the former are costs to process claims, whereas broker commissions are an example for the latter. As detailed in Braun et al. (2019), the incentive compatibility constraint is

$$u_2(s, \theta^i, \Pi^i, b^i) \geq u_2(s, \theta^i, \Pi^j, b^j) \; \forall i, j \in \{t, b\}, i \neq j \quad (8)$$

and the participation constraint is

$$u_2(s, \theta^i, \Pi^i, b^i) \geq u_2(s, \theta^i, 0, 0) \; \forall i \in \{t, b\} \quad (9)$$

5.1.3 Parameters and Parameterization of Model

We followBraun et al. (2019) in their parameterization strategy; for example, we set the real interest rate r to zero. Further, we employ a standard utility function with constant-relative risk aversion

$$u(c) = \frac{c^{1-\sigma}}{1-\sigma}$$

and set the risk aversion parameter σ to 2. There are a series of additional model parameters that we calibrate in a first step. The objective of the calibration is always to match actual data moments. To do so, we rely on various data sources and the stylized facts as presented in Section 4.2. Table 2 lists the main model parameters.

[Insert Table 2 about here]

First step: Calibration of Model Parameters. A first important step is to model the representative health risk score distribution in Figure 4. A beta distribution with $\beta(1.2269; 6.9219)$ approximates this skewed distribution reasonably well. As illustrated in Figures 5 to 7, to keep the data and modeling process tractable, we categorize the continuous health risk score as well as household income and focus on population quintiles. The mean risk scores by the five income quintiles are in Table A7. Following Braun et al. (2019), we assume that their joint distribution follows a Gaussian copula with parameter φ, chosen to match the data points in Table A7.

We use the representative SOEP to extract the wage distribution of those who enter the labor market at the beginning of their work lives between age 25 and 30. We model it as a log normal distribution and normalize Period 1 (Figure 8) to 1. Again, following Braun et al. (2019), we express the consumption floor as a share of *permanent* lifecycle income, which is 0.1258 for Germany.[25] Similarly calculated is the cost of work disability $w_2 - C$ which uninsured individuals incur for 16 years, on average between age 46 and early retirement at age 62, see Figure 8.[26] For the fixed (γ) and variable (λ) administrative costs, we take industry averages of 3% of lifetime and 10% of annual premiums, respectively (Finanzberatung Bierl 2022).

Second step: Matching Simulated Moments. In a second step, we calculate model equilibria and set key parameters to minimize the distance between the actual data moments and the model equivalents. Note that the model has a very high computational intensity and, thus, is not formally estimated, see Braun et al. (2019). One reason for this computational intensity is that the menus of optimal insurance policies that insurers offer must be calculated for each of the 750 different risk groups. The risk groups consist of combinations of health risk (h), income (w), and occupational groups (o).

We simultaneously set parameters minimizing distances between actual data moments and equilibrium model outcomes: the distribution of income uncertainty τ, the work disability risk θ

[25] Permanent income is simply the average gross wage (2019: €47,928, Statistisches Bundesamt (2022)) multiplied by the average contract duration (31.5 years), which is roughly the number of years between signing a contract and retirement. The consumption floor equals the value of the means-tested basic income (2019: €11,868, Bundesagentur für Arbeit (2019)) which people would receive for 16 years between the average age of work disability, 46, and early retirement.

[26] (€47,928-€11,868)×16/(€47,928×31.5)

by health and income, a fraction of good types ψ and the preference parameter β. For example, one target is the 25 work disability probability moments by income and health risk quintiles as in Figure 5.[27] Table A7 shows the actual data moments and model counterparts for health risk by the five income quintiles. As seen, the model produces a very close match between the two.

Intuition of Mechanism of General Equilibrium Model. The model features an optimal contracting framework that includes a means-tested basic income program, private information, and administrative costs. These ingredients are sufficient to replicate the stylized empirical facts of ODI take-up observed in representative data. In particular, they reproduce rising take-up rates in better health and income, although the work disability risk *decreases* in better health and income. Appendix C provides a discussion of possible equilibria and optimal insurance policies for the two risk groups by the insurer.

One main underlying mechanism for low take-up rates is coverage denial, as frequently observed in reality. Insurers decide whether to offer coverage to a risk group after having observed h, w, o. If contracts with reasonable premiums for applicants are unprofitable for the insurer, they deny coverage to some of the 750 risk groups in Period 2. The technical reason for unprofitable contracts are administrative costs, see Chade and Schlee (2020), but also the social safety net, see Braun et al. (2019).

Further, insurers are aware of individuals' optimization problem and that low-income individuals may be better off not paying premiums Π for supplemental insurance that provides little utility, given that they likely qualify for the means-tested basic income program after an income or health shock. However, insurers do not know whether applicants' true work disability risk is high or low, given their observables h, w, o. This is private information to the individual.

5.1.4 Baseline Economy

Panel A of Table 3 shows the 25 ODI take-up moments by income and health quintiles as shown in Figure 6. The columns indicate the five health risk quintiles. The first five rows indicate the income quintiles. The cell in the upper left corner indicates a take-up rate of 25.9% for the bottom income and upper health quintile. Take-up rates increase with higher income quintiles to 45.2% for the richest and healthiest quintile. They fall in bad health to 9.6% for the poorest and sickest quintile, and to 29.1% for the richest and sickest quintile (see also Figure 6).

[27] Following Braun et al. (2019) we assume that the work disability risk is invariant within each of the 25 cells and that applicants know their true risk (θ^t, θ^b)

[Insert Table 3 about here]

Panel B of Table 3 shows the 25 ODI take-up moments by income and health quintiles as produced by the model. As seen, the fit between the empirical and model moments is very close but naturally not perfect. For example, the model is capable of producing a private ODI take-up rate of 23.0% instead of 25.9% for the poorest but healthiest quintile; for the healthiest and richest quintile it is 45.7% instead of 45.2%, and for the poorest and sickest quintile it is 8.3% instead of 9.6%.

5.1.5 Reform Validation

Next, we simulate the reform of the public ODI system and validdate it with our empirically elicited reform effects by health risk quintiles. Figure 9 shows the reform effects by quintiles of the health risk score. It is estimated using our standard bias-correct local polynomial RD regressions (Calonico et al., 2014, 2017, 2018), see equation (2) and Section 4.1,

[Insert Figure 9 about here]

While the average effect in Figure 3 is statistically insignificant, Figure 9 shows a positive and significant reform effect for the healthiest quintile by more than 20 percentage points. However, for the other three health risk quintiles, the point estimates are close to zero but imprecisely estimated. However, for example, we can exclude that the reform increased take-up in quintile 2 by more than 12 percentage points, and in quintile 5 by more than 23 percentage points.

Unfortunately, there exists no pre-reform SAVE data. Consequently, we use the model to simulate the *reverse* reform effect in order to validate it. To do so, we simulate the pre-reform replacement rate for public ODI to accurately assess the costs of an occupational disability in percent of former income. Appendix B illustrates the details of our stylized benefit simulation. Importantly, we assume that the individual starts working at age 25 and earns 60% of the average German wage when entering the labor market. We assume that the wage position would increase linearly to 140% if the individual worked until age 65. Figure B2 (Appendix B) plots the gross replacement rate as a function of the occupational disability age on the x-axis. As seen, it is decreasing from age 25 to 60 because of the linearly increasing wage over the same time period and actuarial deductions. After age 60, when people are closer to the actual retirement age, it sharply increases again. The average gross replacement rate is 18%.

Hence, we simulate the reform effect using the model by letting the costs of an occupational work disability decrease to 82% of the previous wage *for all but the bottom income quintile*.[28] This is important as, pre-reform, public ODI provided not much value to low-income earners, whereas it provided a potentially high value, the larger the income, and the more the income deviates from the consumption floor. Recall that the basic public WDI was not altered through the reform. It provides coverage in case of general work disability, roughly at the same level as the consumption floor in Germany.

The simulation yields a reform effect that is very close to the empirically elicited reform effect for the top health risk quintile, namely 19 percentage points (vs. 23ppt in Figure 9). Further, the reform effect for the other quintiles lies within the confidence intervals in Figure 9.

5.1.6 Policy Simulation

Finally, we use the model to simulate policy counterfactuals. Specifically, we leverage the three main driving forces of the model which are capable of producing the empirically observed income and health gradients in take-up: the social safety net providing a basic consumption floor, private information, and administrative costs.

[Insert Figure 10 about here]

Figure 10 simulates and illustrates the changes in private ODI take-up rates for the three scenarios: full information, no administrative costs, and no basic income. Here, we solely show the results by health risk quintiles, defined as before. In the baseline scenario and for the healthiest three quintiles of the population, average take-up rates are relatively stable—just below 40%—but decrease to 30% and below 20% for quintiles four and five, respectively.

The full information scenario—where insurers would perfectly observe not just $\theta^i_{h,w,o}$ but could also differentiate between the bottom and top tails of the distribution θ^t, θ^b—would shift take-up rates up, almost in a parallel fashion, by around 20 percentage points for most quintiles. The reason is the elimination of frictions and uncertainty, resulting in more targeted offers and policies. Figure 11 shows the mirror image of a higher take-up—falling denial rates across the health distribution. However, besides the question of how a realistic policy option to achieve that goal could look like, we note that we would still observe a clear health gradient in take-up, see Figure 10.

[Insert Figure 11 about here]

[28] Further, we take into account that through several reforms of the German social insurance system, the consumption floor in the re-reform era was more generous. For simplicity, we assume a flat 10% higher consumption floor pre-reform.

By contrast, imagine the German welfare state would not exist and would not provide a hard—and in an international comparison generous—consumption floor. Figure 10 illustrates that the safety net does not matter much for the healthiest 60% of the population. However, absent a public safey net, take-up rates would strongly increase to more than 70% for the sickest two quintiles. Without a consumption floor, those who expect potentially high costs due to a health shock that threatens their work capacity have a much higher demand for and utility from private ODI. As discussed, in the model framework, insurers are aware of individuals' optimization procedure—only the unobserved health type is private information; thus, they are much more likely to offer them (still profitable) coverage absent the safety net, see Figure 11, knowing that they are more likely to accept these offers, given the participation and incentive compatibility constraints. It is worthwhile to emphasize that the scenario without a public safety net produces private ODI coverage gradients that align with the underlying work disability risk (Figure 5), producing a gradient that slopes *upward* instead of downward in bad health.

Finally, a world without administrative costs would result in the highest coverage rates of all scenarios—around or above 80% for most health quintiles—except for the sickest quintile whose coverage rates would, however, still increase to 70%. Recall that administrative costs are substantial in this individual market—fixed costs like broker commissions amount to an estimated three percent of lifetime premium payments over 31.5 years, in addition to ten percent reoccurring annual administrative costs. All these costs drive up the price for insurance coverage. They lead to high denial rates and applicants who are unwilling to pay these high premiums—given the consumption floor provided by the welfare state, their expected lifetime income (including income shocks) as well as expected health shocks during their main working age. Consequently, eliminating administrative costs would strongly reduce denial rates across the health distribution (Figure 11).

Appendix Table A8 shows the full simulated take-up rates, not only by health, but also by income quintiles. Here we see a more nuanced picture. For example, without private information, we would still see income gradients in coverage for the healthiest (33.0% to 46.1%) and second healthiest quintiles (29.2% to 41.1%); but much flatter ones. In a world without safety net, the income gradients in private ODI take-up would flip (for all except the healthiest quintile) and align—as expected—with the strong income gradient in work disability risk as shown in Figure 5.

Table 4 shows overall take-up rates, the share of insured costs, and loading factors for the baseline economy and various policy scenarios. This time, the table differentiates between good and bad risks, that is, θ^b and θ^t. As shown in Table 2, 73% fall into the bottom of the disability risk distribution. Thus, for about three-quarters of the population, the absence of admin costs and a safety net would not just increase take-up rates, but also the share of the insured work disability risk from 84% to more than 90%.[29] However, if insurers could identify the true types with certainty absent information asymmetries, optimal policies would entail a smaller share of insured risk for the good types, but a substantially higher share of insured risk for the bad types. Moreover, the share of insured risk would be relatively even across health risks, without much of a gradient, see Appendix Figure A13. The reason is that insurers would offer lower and higher coverage contracts that are better tailored for the lifecycle optimization problem of the good and bad risks, respectively, given their expected income and health shocks and the guaranteed consumption floor by the government.

As a very last outcome, the model produces predictions for the loading factor, which is defined as one minus the ratio of the expected value of benefits to premia; thus a load of zero would indicate an actuarially fair contract. In the baseline scenario, insurers load contracts for bad risks substantially, whereas the loads for good risks are negative. The loads decrease without admin costs and private information for the bad risks. They become positive for the good types without a social safety net and private information. The reason is how insurers calculate the optimal contracts that they decide to offer to each type: As discussed, denial rates decrease under both scenarios (but without a safety net, only for the sick, Figure 11). In a world without safety net, while overall take-up rates would remain stable for the majority of the population (and increase for the sick), insurers would increase both, the insured risk but also the loads of the policies offered (Figure 10). By contrast, in a world without information frictions in which they could identify good risks with certainty, overall take-up for good risks would remain relatively stable but their policies would not just cover a substantially smaller share of the risk (which decreases premiums substantially) but also carry higher loads (which increases premiums substantially) compared to the baseline scenario. Interestingly, the absence of information frictions would benefit the bad risks, as their take-up rates would almost double and their contracts insure a much greater share of the work disability risk, plus the loads would decrease.

Summary. When holistically assessing the counterfactual simulations in conjunction with their implied policy alternatives, a relatively clear picture emerges. First, as Germany has just increased the means-tested consumption floor by 12%, the asset eligibility thresholds by 50% (Deutscher Bundestag, 2022), WDI benefits in 2014 and 2019 (see footnote 9), and also reduced possibilities to sanction those who receive cash benefits but are unwilling to cooperate with caseworkers,

[29]This share refers to the drop in income in case of work disability.

political reforms to lower the consumption floor are very likely politically infeasible. In a 2018 representative poll, 53% of Germans found the benefit levels inappropriately low (YouGov, 2018). Second, while eliminating frictions and asymmetric information would be desirable, given the relatively unregulated market and wide possibilities to risk-rate policies, there is no obvious policy to address this issue. Policymakers could allow genetic sampling which, however, would not eliminate private information and would be highly controversial, especially in Germany with its history. Moreover, this policy option would not eliminate the strong health and income gradients in coverage. The final option, reducing administrative costs, emerges clearly as the most desirable and feasible of all policy alternatives. Our findings show: Not only is the potential to increase take-up the largest, also the health gradient in coverage would be substantially reduced. Most importantly, there exist clear regulatory tools with bipartisan support to implement such a policy, for example, a cap on commission fees or minimum benefit ratios. Benefit loss ratios would cap the ratio of benefit payouts to administrative costs, e.g. spending on marketing by insurers.

6 Discussion and Conclusion

This paper studies a structural reform of the disability insurance market in Germany, both empirically and theoretically using reduced-form approaches and a general equilibrium model. The reform cut access to public occupational disability insurance (ODI) effective 2001 for cohorts born after 1960. We call these cohorts "notch cohorts."

Unlike older cohorts and during the generous post-WWII German welfare state, the notch cohorts could no longer apply for public benefits when a health shock prevented them from working in their previous *occupation*. This paper first studies the first-stage effects on public DI inflows. Then it studies interaction effects with the biggest private individual ODI market in the world. However, this private individual market is relatively unregulated. Guaranteed issue does not exist, premiums are risk rated and coverage denials common. Applicants purchase policies at an average age of 32, after having entered the labor market and when settling down and starting a family. They keep their policies for an average of 31 years, covering the crucial time period of their work lives until early statutory retirement is possible.

While we find that the reform reduced public DI inflows by 35% in the long-run, we do not find much evidence that the notch cohorts purchased private ODI policies at much higher rates than the non-notch cohorts. We can exclude increases of 12 percentage points with 95%

probability at the population level. Standard welfare models that infer insurance values from demand elasticities and take-up may imply that people don't value this type of supplemental insurance sufficiently. And that the reform increased welfare.

However, the small interaction effects are just an equilibrium outcome. Hence, we employ a general equilibrium model and tailor it to fit the German case, given the regulatory framework of the market. The model features three main driving forces: (i) the German means-tested basic income program which provides a guaranteed consumption floor to all residents of about 100% FPL; (ii) private information, and (iii) administrative costs. We show that the model and these three driving forces are powerful enough to explain stylized empirical pattern in the private ODI market: (1) despite a high lifecycle work disability risk, take-up rates are below 50% across all health and income groups. We find at best modest interaction effects with public DI. Further, we find (2) strong take-up gradients that increase in good health and income, (3) inversely related gradients in the lifecycle work disability risk that decrease in good health and income.

Our simulations suggest that policymakers could increase take-up rates substantially by either (A) streamlining the means-tested basic income program (which would be politically difficult and may have unintended consequences), or (B) allowing insurers to collect even more data to further reduce private information, e.g. via genetic samples (politically even more difficult and with potentially more severe unintended consequences), or (C) implementing market-based reforms that would lower the high administrative costs in this market. For example, it is very common that broker commissions amount to several monthly ODI premiums, where a monthly premium can be as high as several hundred dollars for high risk groups. Concrete policy proposals could suggest to cap or even ban such high commission fees. Alternative policy proposals could limit "benefit loss ratios" akin to the regulation of U.S. health insurers through the Affordable Care Act that imposed medical loss ratios. In contrast to the other measures, targeting administrative costs has the highest potential to increase take-up. Further, it could substantially reduce the health gradient in take-up, even under risk rating. Our simulations predict denial rates to fall across the entire distribution of applicants' health risks; simultaneously, the disability risk covered by the private ODI policies would increase, and loads decrease, for both high and low risk types.

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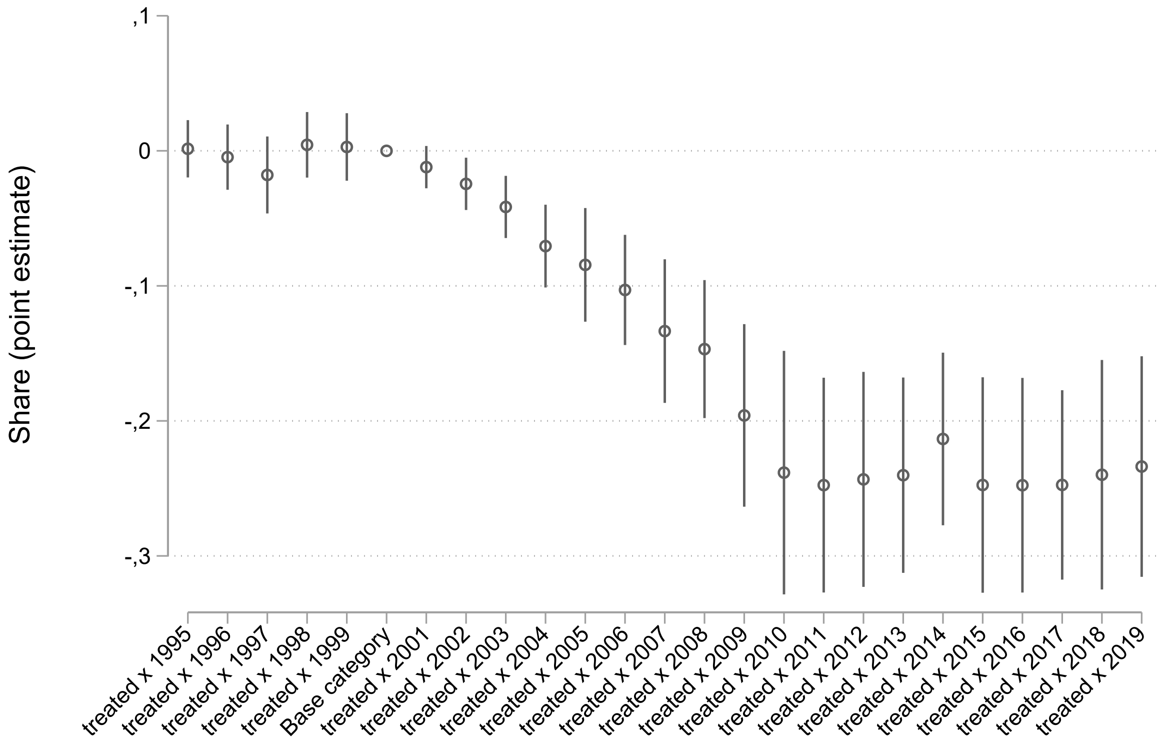
Figure 1: Effect of 2001 Reform on Public DI Inflows Using Administrative Data



Source: Administrative SPI data on new public DI recipients by cohort and year. Notch cohorts are those born after 1960 and the treatment group; grandfathered cohorts are those born before 1961 and the control group. Figure plots $\beta D_c \times T_t$ estimates from equation 1 but with the post-reform indicator T_t replaced by a series of year dummies where 2000 is the base year.

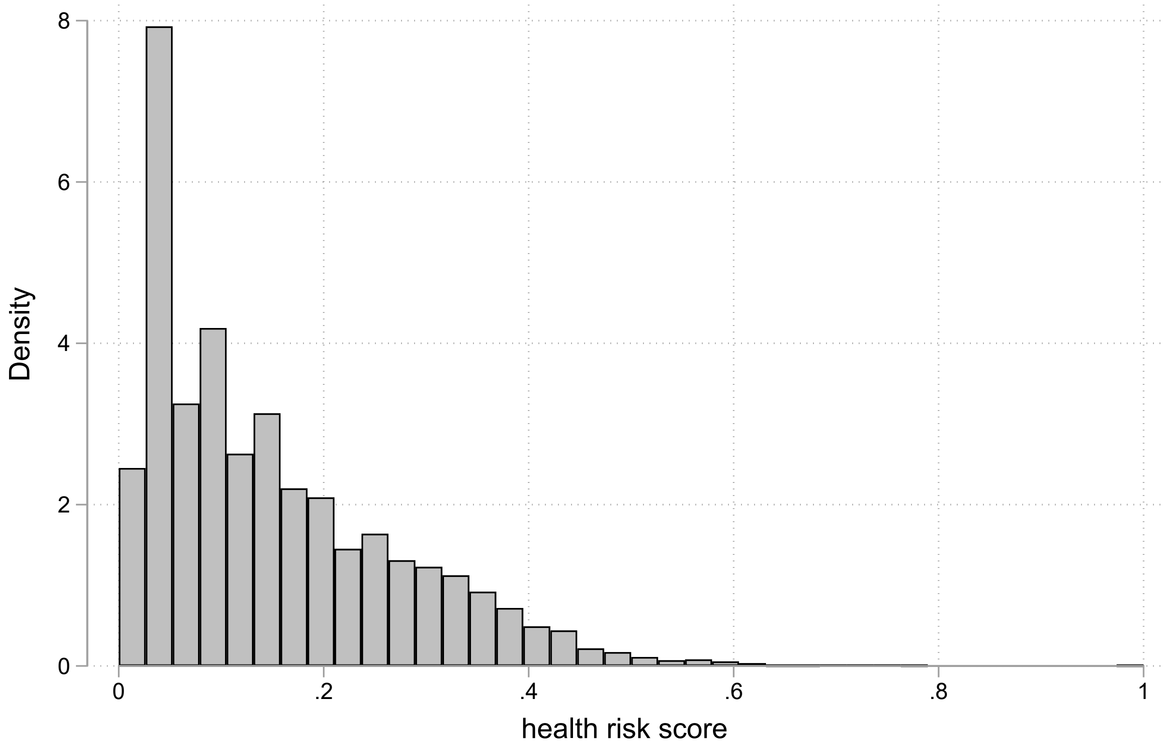
Figure 2: Effect of 2001 Reform on Public DI Case Loads Using Representative SOEP Data

Source: SOEP v.33 – 95% sample. Sample is restricted to post-reform years. The figure is from one RD model similar to equation 2, estimated using quadratic trends in the running variable $f(z_i - c) = z_i - c$ to allow for different slopes before and after the cutoff. Robustness checks show results for an alternative *PublicDI II* measure and the pre-reform period (Figure A5), vary the bandwidth (Figure A6), study the smoothness of covariates (Figure A7), carry out density plots of running variables (Figure A8), and vary polynomials as well as carry out donut RDs (Figure A8).

Figure 3: Effect of 2001 Reform on Private ODI Policies Using Representative SAVE Data

Source: SAVE data 2001-2010. The figure shows the raw nonparametric means of private ODI coverage by birth year, overlaid with separate linear trends before and after the cutoff. Other robustness checks vary the sample (Figure A8), vary the bandwidth (Figure A10, Calonico et al. 2020), study the smoothness of covariates (Figure A11), carry out density plots of the running variable (Figure A12, McCrary 2008), and vary polynomials as well as run donut RDs (Figure A8).

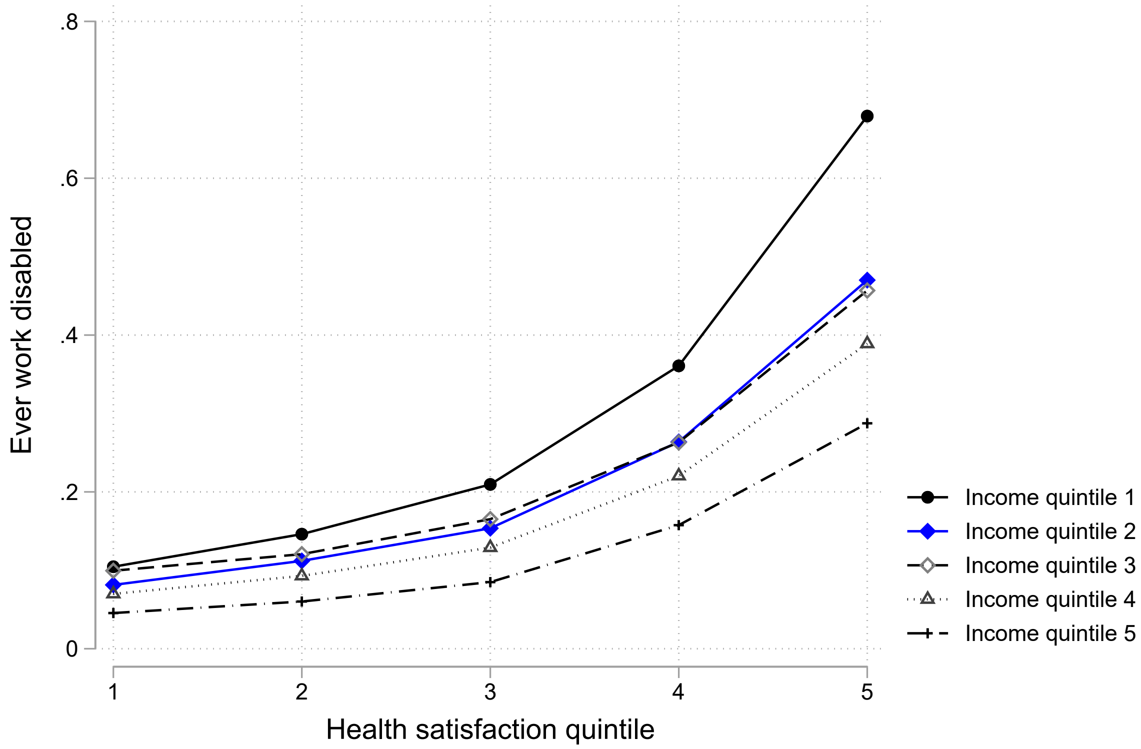
Figure 4: Distribution of Health Risk Score



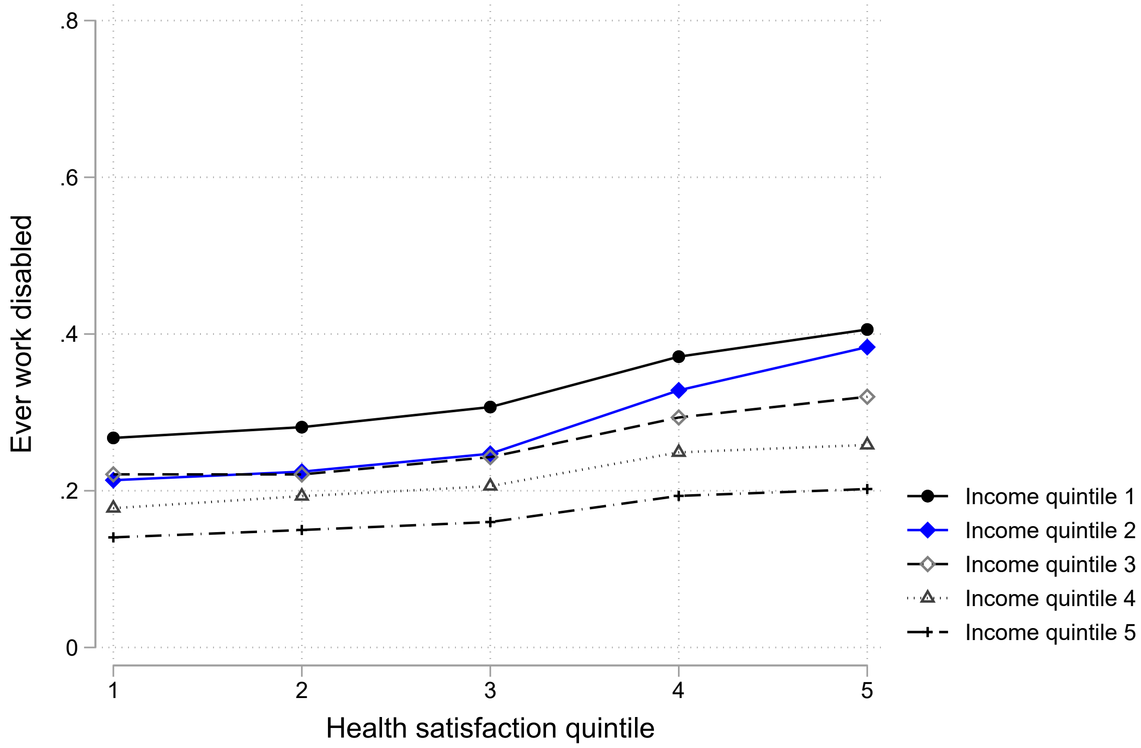
Source: SAVE data 2001-2010. Health risk score is produced using principal component analysis and subjective as well as objective health measures from SAVE.

Figure 5: Lifecycle Risk of Work Disability by Income and Health Risk Score

(a) unconditional

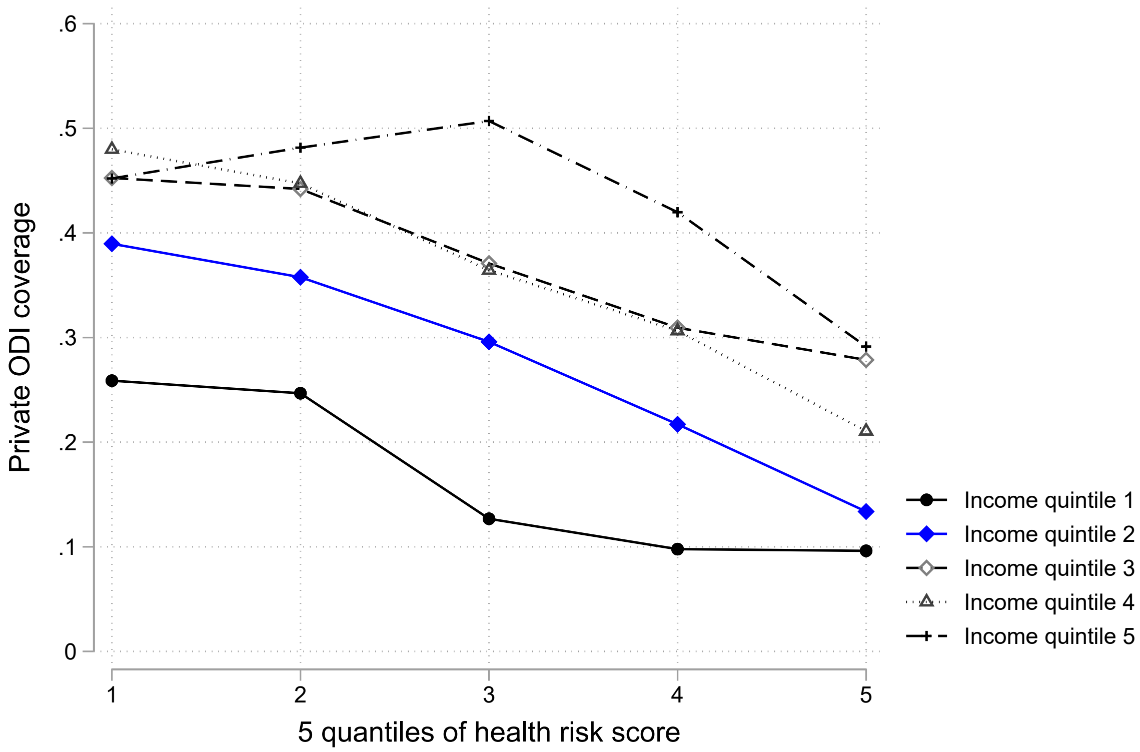


(b) conditional on job, education, socio-demographics



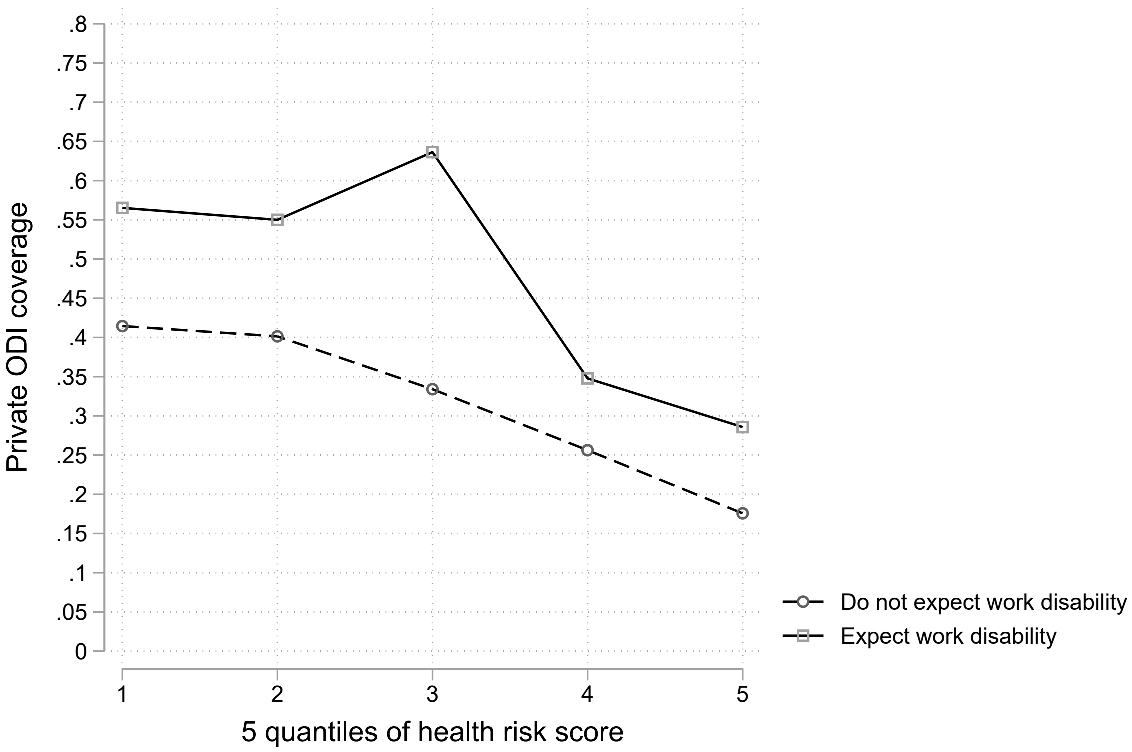
Source: SOEP v.33 – 95% sample. Figure 5a plots the unconditional risk of a severe health limitation over the working ages by the health satisfaction quintiles and the five net household income quintiles. Figure 5b first regresses the lifecycle risk of severe health limitations on socio-demographics, job and educational characteristics, predicts the risk at the individual level and then plots this conditional risk by the health satisfaction quintiles and the five net household income quintiles.

Figure 6: Take-Up of Private ODI Policies by Health Risk and Income



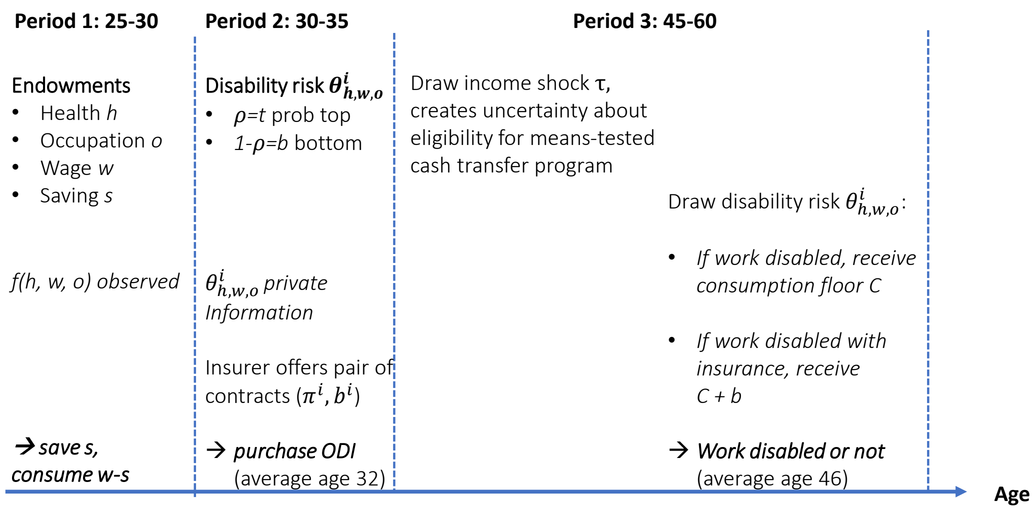
Source: SAVE data 2001-2010. Figure plots take-up rates of private ODI policies against the quintiles of the health risk score in Figure 4 and stratifies these curves by the five net household income quintiles.

Figure 7: Take-Up by Health Risk and Private Information



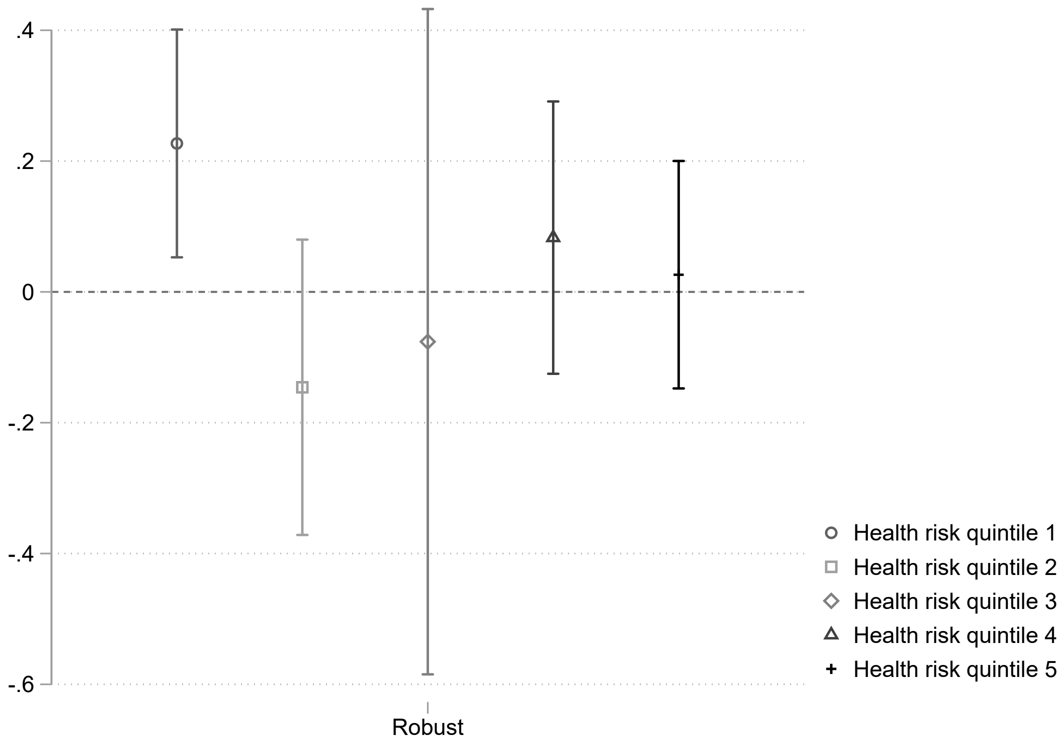
Source: SAVE data 2001-2010. Figure plots take-up rates of private ODI policies against the quintiles of the health risk score in Figure 4 and stratifies these curves by expected retirement before age 60. The latter information is directly elicited in the SAVE survey and proxies expected work disability.

Figure 8: Illustration of Lifecycle Time Periods in Baseline Model



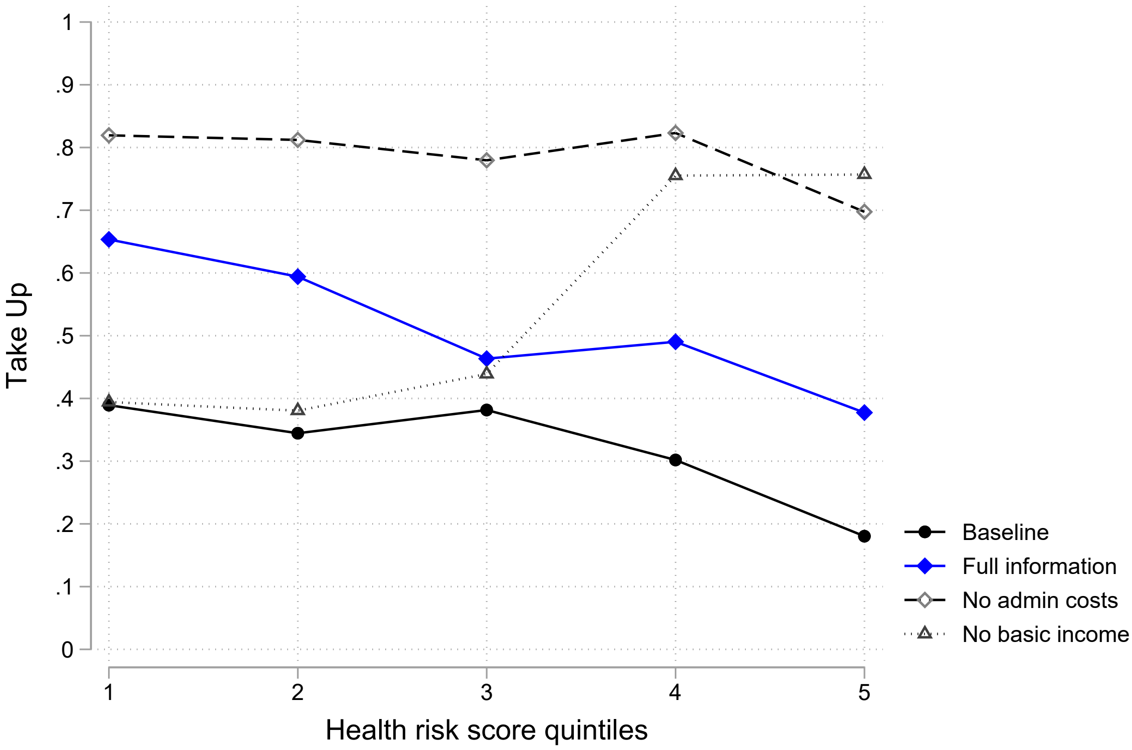
Source: The figure represents a simplification of the lifecycle decision-making process of a variant of the model by Braun et al. (2019). For more details, please see main text.

Figure 9: Effect of 2001 Reform on Private ODI Policies by Health Risk Quintile



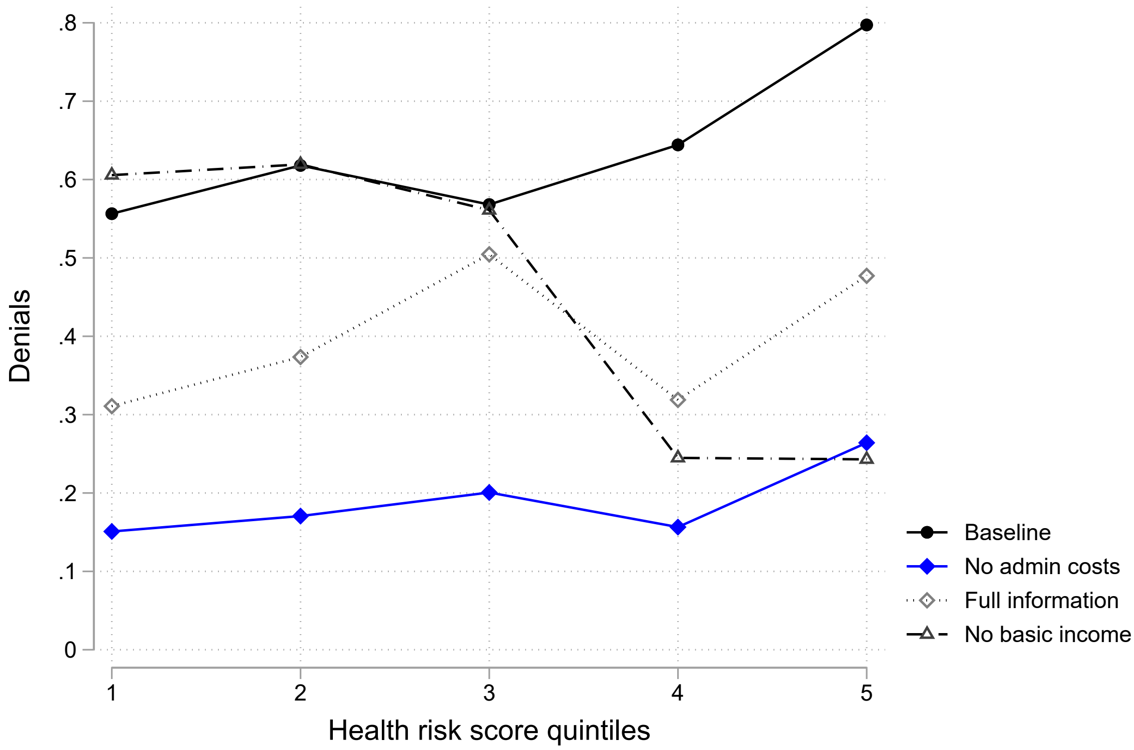
Source: SAVE data 2001-2010. The figure shows the 2001 reform effect by health risk score quintiles using our standard RD models similar to equation (2), estimated using local polynomial regressions with linear polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). See main text for more details.

Figure 10: Take-Up Rates by Health Risk Score: Baseline vs. Policy Simulations



Source: The solid black line represents the baseline private ODI take-up rates by the quintiles of the health risk score in Figure 4. The other lines show take-up rates for alternative policy simulations by health risk quintiles using the general equilibrium model (see Section 5).

Figure 11: Denial Rates by Health Risk Score: Baseline vs. Policy Simulations



Source: The solid black line represents the baseline denial rates by the quintiles of the health risk score in Figure 4. The other lines show take-up rates for alternative policy simulations by health risk quintiles using the general equilibrium model (see Section 5).

Table 1: Effect on Private ODI Coverage Using Representative SAVE Data

	(1) Full sample	(2) SPI insured	(3) No kids	(4) One-person HH
Conventional	-0.045	-0.053	-0.017	0.021
	(0.0348)	(0.0443)	(0.0494)	(0.0691)
Bias-corrected	-0.048	-0.051	0.041	0.029
	(0.0348)	(0.0443)	(0.0494)	(0.0691)
Robust	-0.048	-0.051	0.041	0.029
	(0.0417)	(0.0509)	(0.0572)	(0.0794)
Year FE	yes	yes	yes	yes
State FE	yes	yes	yes	yes
Age + gender	yes	yes	yes	yes
Work + education	no	no	no	no
Socio-dems	no	no	no	no
Conventional	-0.057	-0.075**	-0.060	-0.034
	(0.0464)	(0.0351)	(0.0506)	(0.0671)
Bias-corrected	-0.059	-0.100***	-0.052	-0.010
	(0.0464)	(0.0351)	(0.0506)	(0.0671)
Robust	-0.059	-0.100**	-0.052	-0.010
	(0.0536)	(0.0421)	(0.0596)	(0.0760)
Year FE	yes	yes	yes	yes
State FE	yes	yes	yes	yes
Age + gender	yes	yes	yes	yes
Work + education	yes	yes	yes	yes
Socio-dems	yes	yes	yes	yes
Observations	12822	9,580	6,236	2,281

Source: SOEP v.33 – 95% sample. Table reports estimates for RD models similar to equation (2), estimated using local polynomial regressions with linear polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary the bandwidth (Figure A9, Calonico et al. 2020), study discontinuities in covariates (Figure A10), carry out density plots of the running variable (Figure A11, McCrary 2008), and vary polynomials as well as run donut RDs (Figure A12)

Table 2: Model Parameters

Interest rate	r	0
Risk aversion	σ	2
Health risk distribution	f	$\beta(1.2269; 6.9219)$
Copula parameter	φ	-0.29
Period 1 wage distribution	w	ln(w) $\sim$ N(-0.32, 0.64)
Basic cash income consumption floor	c	0.1258
Work disability costs ($w - C$)	m	0.3822
Insurer's variable costs	λ	1.1
Insurer's fixed costs	γ	1.03
Preference discount factor	β	0.94
Income shock distribution	τ	1-τ truncated log normal
τ bounds	μ_τ	[-2;0.5]
Fraction good types	ψ	0.73

Source: SAVE for frailty distribution, SOEP for young endowment distribution, demand shock distribution, τ , own calculations and various sources for insurer administrative costs and the welfare consumption floor (Bundesagentur für Arbeit, 2019).

Table 3: Private ODI Take-Up Rates by Income and Health Quintiles: Data and Model Fit

	Health Risk Quintile				
Income Quintile	**Q1**	**Q2**	**Q3**	**Q4**	**Q5**
Panel A: Data					
Q1	0.2588	0.2468	0.1268	0.0978	0.0962
...	0.3896	0.3577	0.2959	0.2171	0.1337
...	0.4525	0.4420	0.3709	0.3094	0.2786
...	0.4799	0.4474	0.3643	0.3064	0.2105
Q5	0.4521	0.4815	0.5069	0.4198	0.2914
Panel B: Model					
Q1	0.2302	0.2323	0.1158	0.1289	0.0830
...	0.3938	0.3808	0.2783	0.2145	0.1356
...	0.4628	0.4646	0.3725	0.3139	0.2935
...	0.4495	0.4719	0.3539	0.3112	0.2024
Q5	0.4570	0.5016	0.5379	0.4323	0.2590

Table shows private ODI take-up rates by Health Risk (columns) and Income Quintiles (Rows). Q1 is the healthiest and poorest quintile, whereas Q5 is the sickest and richest quintile. Panel A shows the raw data from SAVE and Panel B show the private ODI take-up rates as produced by the general equilibrium model.

Table 4: Take-Up, Loading, and Risk Insured: Baseline vs. Policy Simulations

	Baseline	**No Admin Costs**	**No Basic Income**	**No Private Information**
Panel A: Total				
Take-up rate	0.3196	0.7864	0.5452	0.5157
Share of costs insured	0.7520	0.8019	0.8978	0.8629
Loading	0.3195	0.3026	0.6718	0.5772
Panel B: Bad risks				
Take-up rate	0.3632	0.8114	0.5452	0.3399
Share of costs insured	0.8271	0.9612	0.9448	0.4144
Loading	-0.6785	-0.7242	0.1345	0.1835
Panel B: Good risks				
Take-up rate	0.3034	0.7771	0.5452	0.5808
Share of costs insured	0.7187	0.7404	0.8804	0.9599
Loading	0.7613	0.6991	0.8705	0.6625

Table shows private ODI take-up rates, share of costs insured and loading factors by scenarios and types.

Appendix A

Figure A1: Reforms and DI Recipiency Rate as a Share of the Working Population

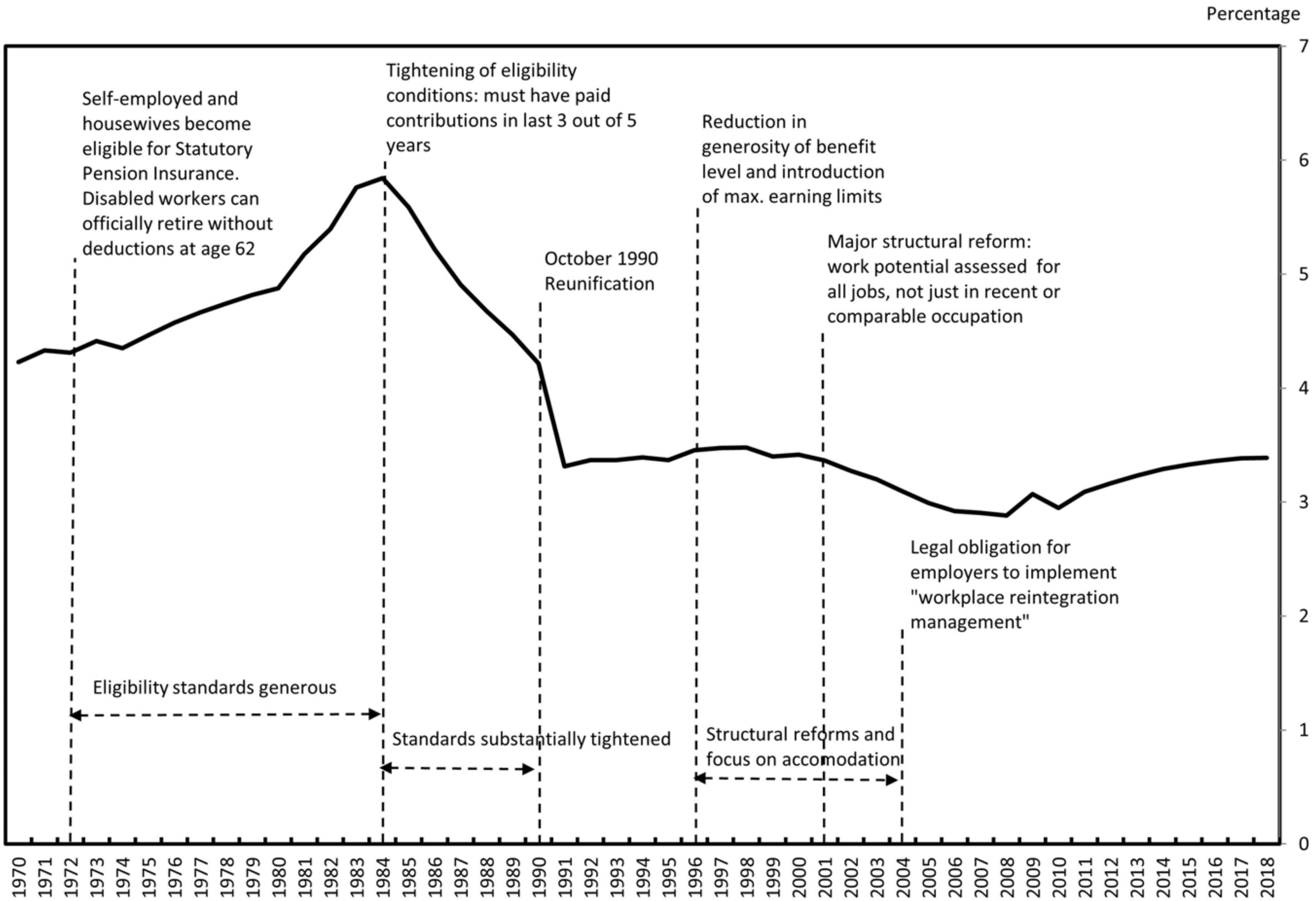
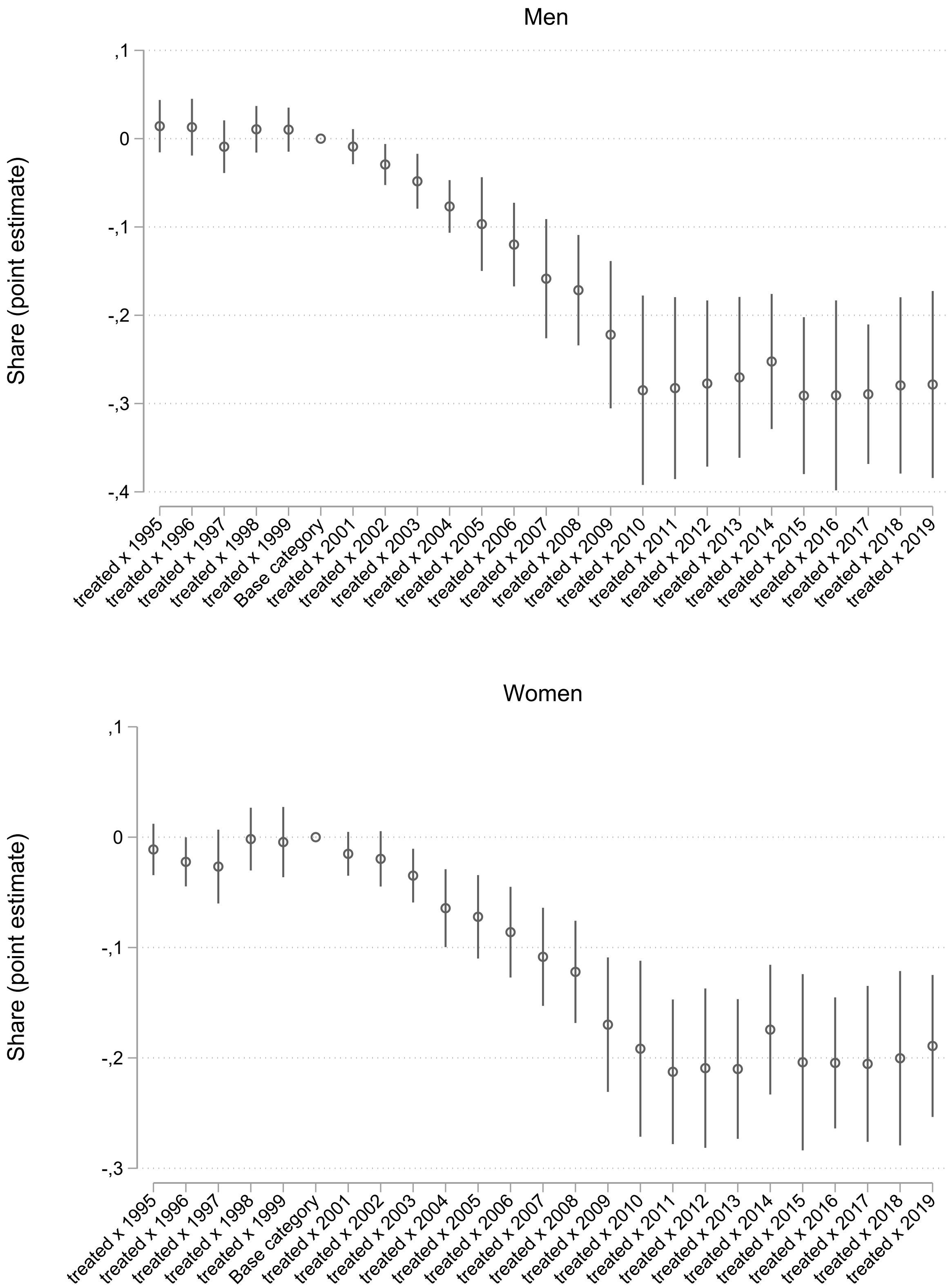
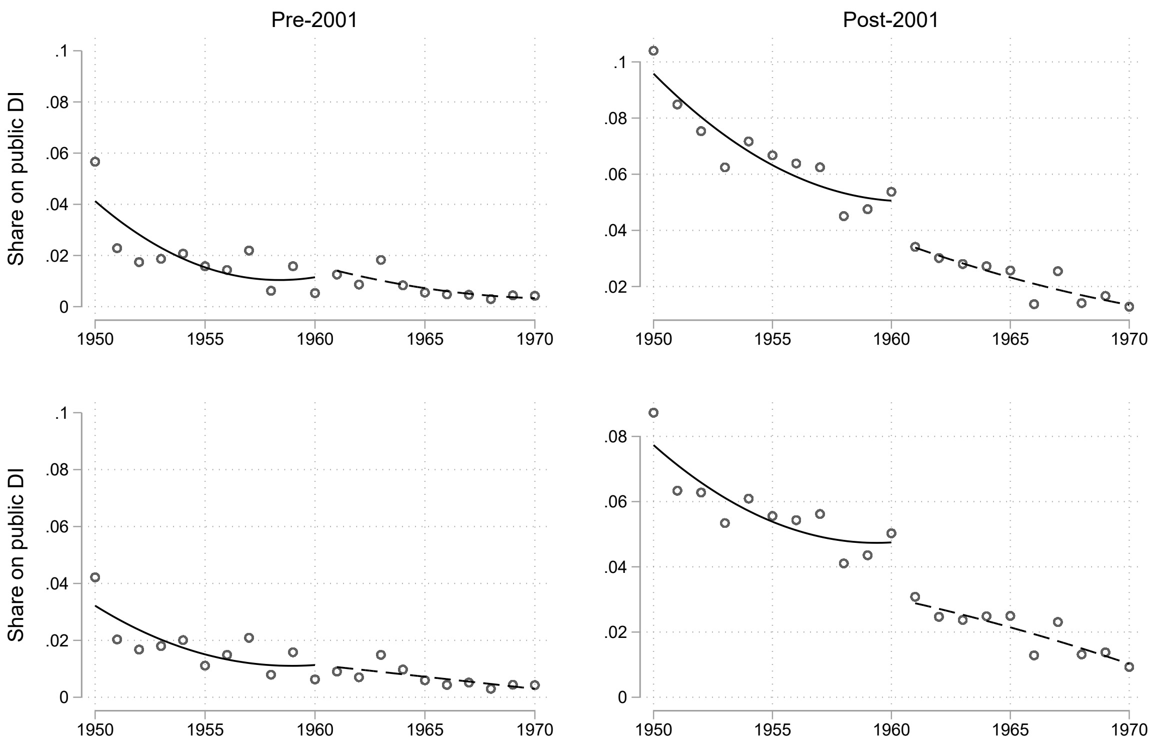


Figure A2: Effect of 2001 Reform on Public DI Inflows by Gender



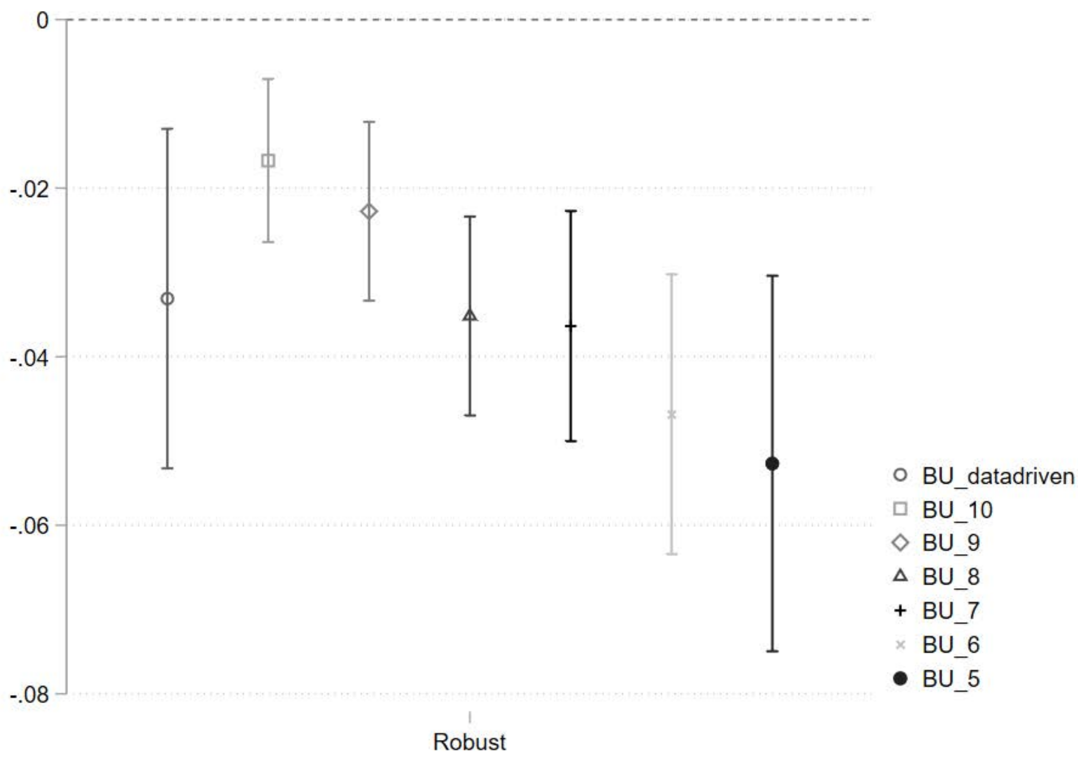
Source: Administrative SPI data on new public DI recipients by cohort and year. Notch cohorts are those born after 1960 and the treatment group; grandfathered cohorts are those born before 1961 and the control group. Figure plots $\beta D_c \times T_t$ estimates from equation 1 but with the post-reform indicator T_t replaced by a series of year dummies where 2000 is the base year.

Figure A3: Effect of 2001 Reform on Public DI Using Representative SOEP Data (II)



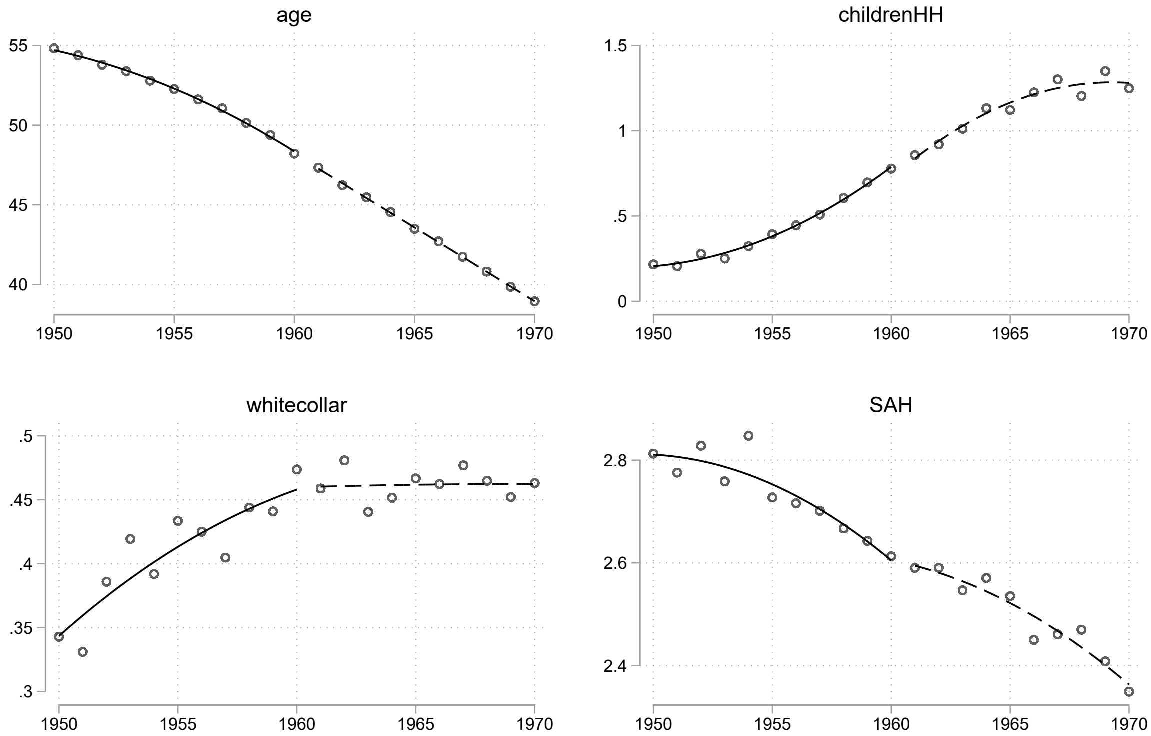
Source: SOEP v.33 – 95% sample. Left column shows pre-reform and right column shows post-reform years. The first row shows *Public DI I* and the second row shows *Public DI II.* All figures show the raw nonparametric means of public disability receipt by birth year, overlaid with separate quadratic trends before and after the cutoff. Other robustness checks vary the bandwidth (Figure A4, Calonico et al. (2020), study the smoothness of covariates (Figure A5), carry out density plots of the running variable (Figure A6, McCrary (2008), and vary polynomials as well as run donut RDs (Figure A7).

Figure A4: Effect of 2001 Reform—Local Polynominal RD Regressions Varying Bandwidth



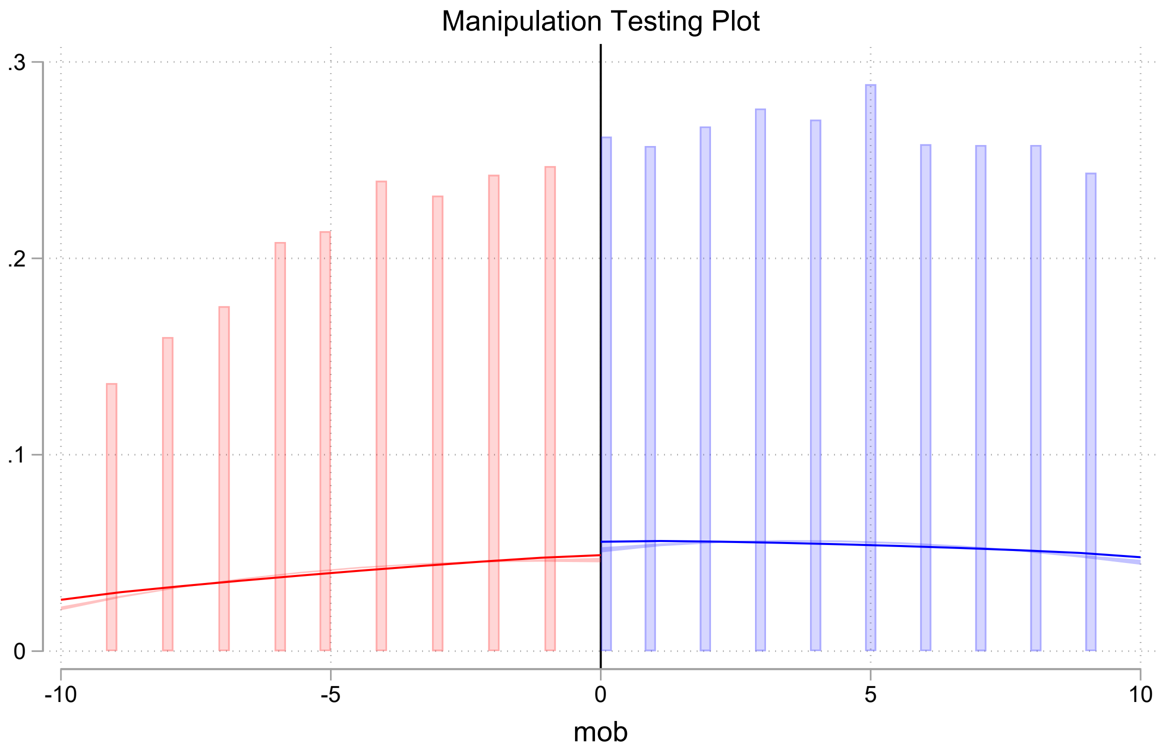
Source: SOEP v.33 – 95% sample. The figures show point estimates of robustness checks varying the bandwidths of RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary the sample and indicator (Figure A3), vary the bandwidth (Figure A4, Calonico et al. 2020), study the smoothness of covariates (Figure A5), carry out density plots of running variables (Figure A6, McCrary 2008) and vary polynomials as well as run donut RDs (Figure A7).

Figure A5: Effect of 2001 Reform—Smoothness of Covariates



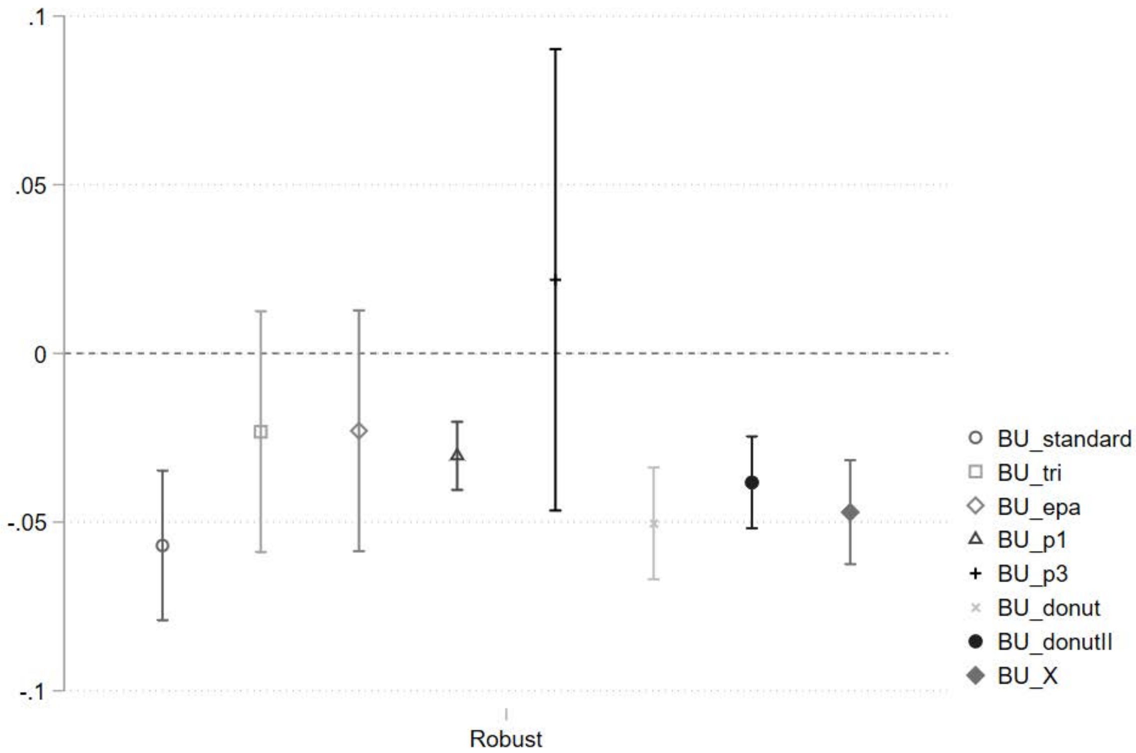
Source: SOEP v.33 – 95% sample. The figures show the raw nonparametric means of covariates as indicated, by birth year, overlaid with separate quadratic trends before and after the cutoff. Other robustness checks vary the sample and indicator (Figure A3), vary the bandwidth (Figure A4, Calonico et al. 2020), carry out density plots of running variables (Figure A6, McCrary 2008), and vary polynomials as well as carry out donut RDs (Figure A7).

Figure A6: Effect of 2001 Reform—Density Plot



Source: SOEP v.33 – 95% sample. The figures shows a density plot of the running variable for RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary the sample and indicator (Figure A3), vary the bandwidth (Figure A4, Calonico et al. 2020), study the smoothness of covariates (Figure A5), carry out density plots of running variables (Figure A6, McCrary 2008) and vary polynomials as well as run donut RDs (Figure A7).

Figure A7: Effect of 2001 Reform—Local Polynominal RD—-Further Robustness

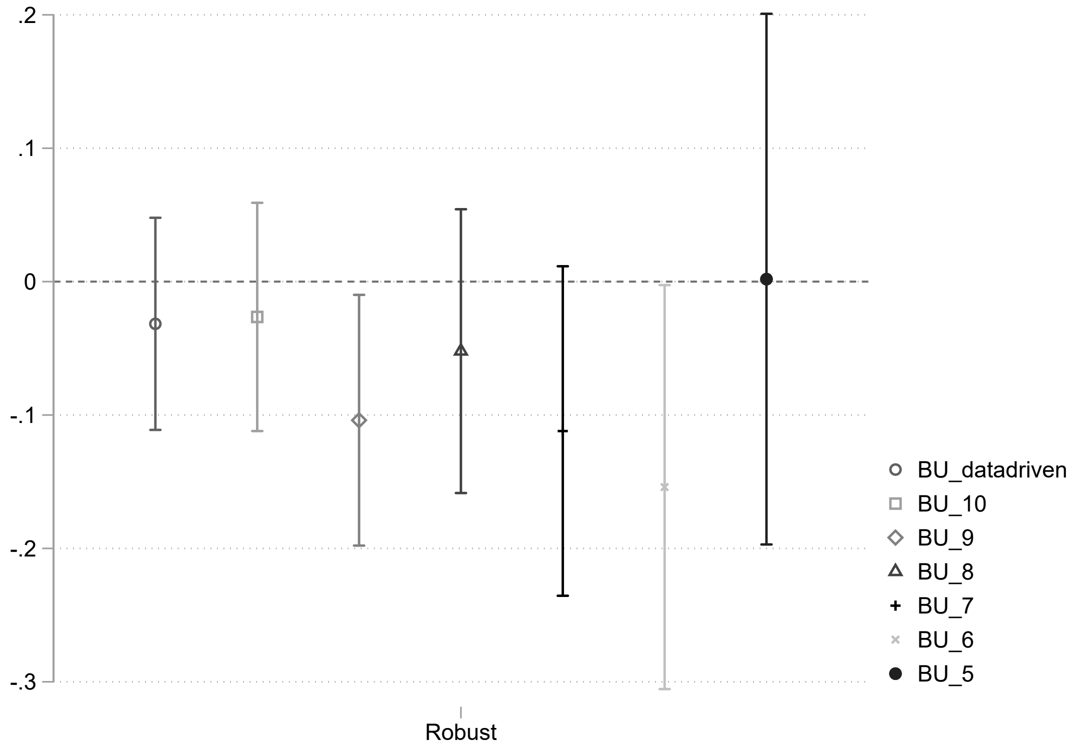


Source: SOEP v.33 – 95% sample. The figure shows the point estimates of a robustness check varying the order of the polynomials, varying weights, adding covariates,m and running donut RD models similar to equation (2), estimated using local polynomial regressions (Calonico et al. 2014, 2017, 2018, 2019). Other robustness checks vary the sample and indicator (Figure A3), vary the bandwidth (Figure A4, Calonico et al. 2020), study the smoothness of covariates (Figure A5), carry out density plots of running variables (Figure A6, McCrary 2008) as carry out donut RDs (Figure A7).

Figure A8: Effect on Private ODI Coverage Using Representative SAVE Data (II)

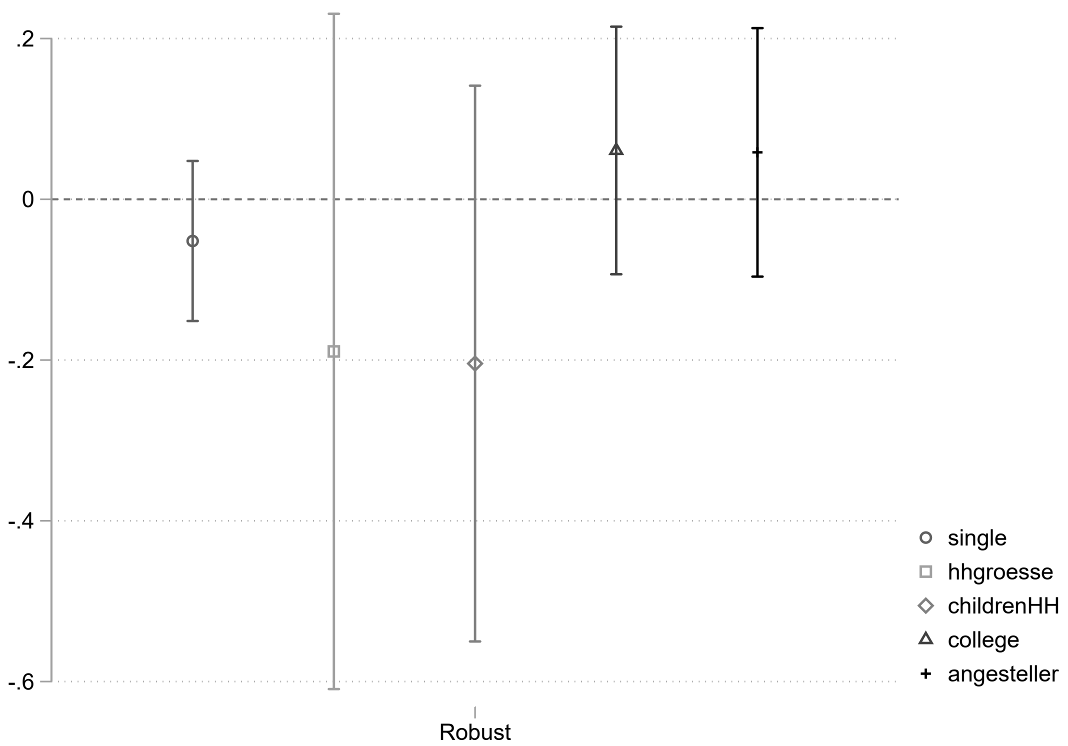
Source: SAVE data 2001-2010. The figures show the raw nonparametric means of private ODI coverage by birth year, overlaid with separate linear trends before and after the cutoff. The upper left graph is the default Figure (3), the upper right figure focuses on those eligible for Public DI, the bottom left focuses on the childless, and the bottom right on one-person households. Other robustness checks vary the bandwidth (Figure A9, Calonico et al. 2020), study discontinuities in covariates (Figure A10), carry out density plots of the running variable (Figure A11, McCrary 2008), and vary polynomials as well as run donut RDs (Figure A12).

Figure A9: Effect on Private ODI Coverage—Local Polynominal RD Varying Bandwidth



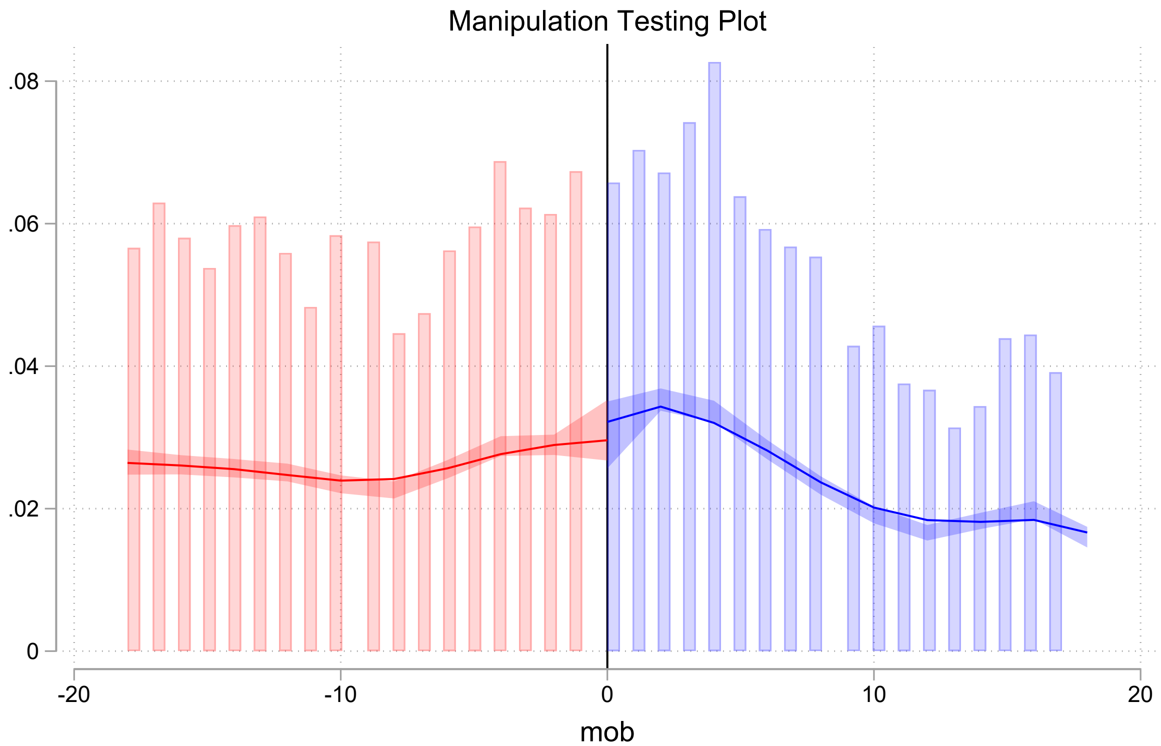
Source: SAVE data 2001-2010. The figures show point estimates of robustness checks varying the bandwidths of RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary the sample (Figure A8), study discontinuities in covariates (Figure A10), carry out density plots of running variables (Figure A11, McCrary 2008) and vary polynomials as well as run donut RDs (Figure A12).

Figure A10: Effect on Private ODI Coverage—Discontinuities in Covariates



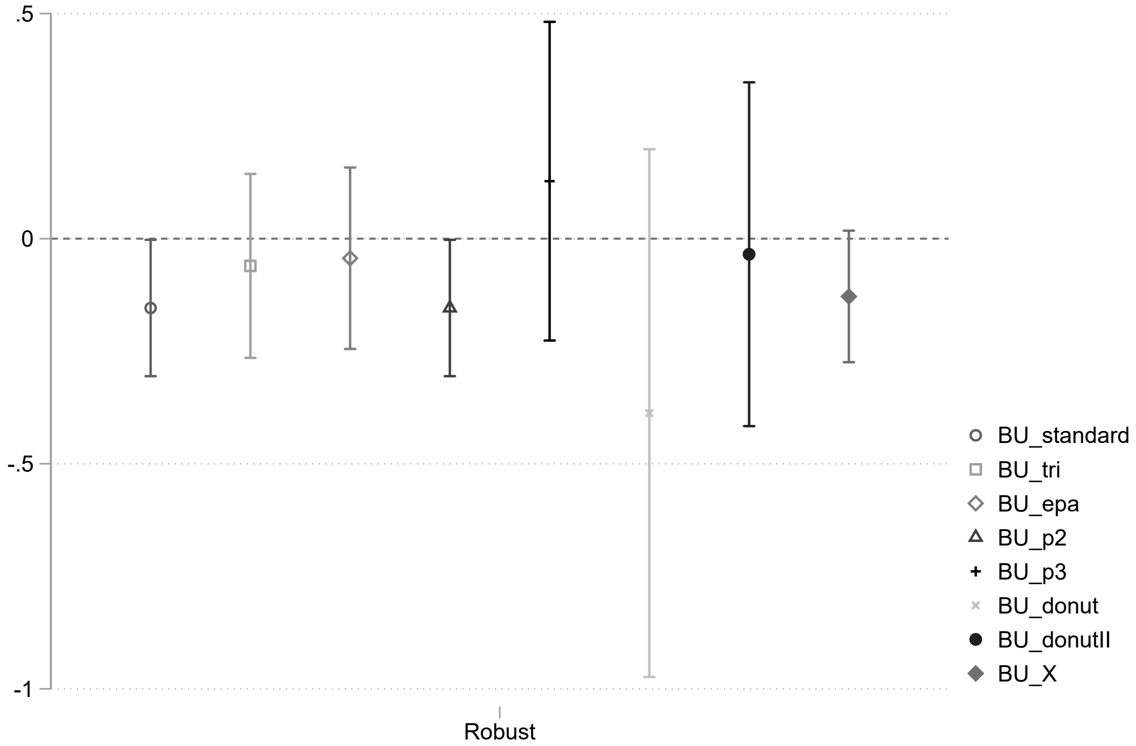
Source: SAVE data 2001-2010. The figures show point estimates of robustness checks testing for discontinuities in covariates using RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary the sample (Figure A8), vary the bandwidth (Figure A9, Calonico et al. 2020), carry out density plots of running variables (Figure A11, McCrary 2008), and vary polynomials as well as carry out donut RDs (Figure A12).

Figure A11: Effect on Private ODI Coverage—Density Plot



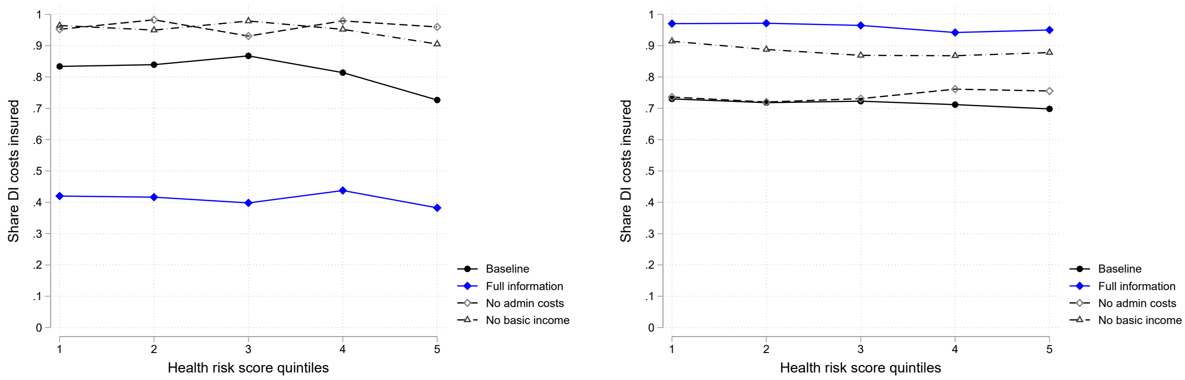
Source: : SAVE data 2001-2010. The figures shows a density plot of the running variable for RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary the sample (Figure A8), vary the bandwidth (Figure A9, Calonico et al. 2020), study discontinuities in covariates (Figure A10), carry out density plots of running variables (Figure A11, McCrary 2008) and vary polynomials as well as run donut RDs (Figure A12).

Figure A12: Effect of 2001 Reform: Local Polynominal RD—-Further Robustness



Source: SAVE data 2001-2010. The figure shows the point estimates of a robustness check varying the order of the polynomials, varying weights, adding covariates, and running donut RD models similar to equation (2), estimated using local polynomial regressions (Calonico et al. 2014, 2017, 2018, 2019). Other robustness checks vary the sample (Figure A8), vary the bandwidth (Figure A9, Calonico et al. 2020), study discontinuities in covariates (Figure A10), and carry out density plots of running variables (Figure A11, McCrary 2008).

Figure A13: Share of Insured Risk by Good and Bad Risk Type: Baseline vs. Policy Simulations



(a) good risk type **(b)** bad risk type

Source: The solid black line represents the share of insured work disability risk by the quintiles of the health risk index in Figure 4. The other lines show take-up rates for alternative policy simulations by health risk quintiles using the general equilibrium model (see Section 5). Subfigure (a) shows the results for the good risk types and subfigure (b) for the bad risk types.

Table A1: Impact on Public DI Inflows Using Administrative SPI Data

Panel A. All	(1)	(2)	(3)	(4)	(5)
$D_c \times T_t$	-0.0907***	-0.0907***	-0.0907***	-0.144***	-0.0514***
	(0.0293)	(0.0219)	(0.0184)	(0.00992)	(0.0105)
D_c	0.364***	0.485***	0.485***	0.762***	0.774***
	(0.0199)	(0.0344)	(0.0289)	(0.0192)	(0.0204)
T_t	-0.159***	-0.266***	-0.266***	-0.397***	-0.0782***
	(0.0255)	(0.0290)	(0.0243)	(0.0137)	(0.0101)
N	1,300	1,300	1,300	1,164	388
Control group mean	0.61	0.61	0.61	0.58	0.50
Panel B. Men					
$D_c \times T_t$	-0.127***	-0.127***	-0.127***	-0.174***	-0.0649**
	(0.0224)	(0.0230)	(0.0231)	(0.0275)	(0.0170)
N	650	650	650	582	194
Control group mean	0.65	0.65	0.65	0.61	0.52
Panel C. Women					
$D_c \times T_t$	-0.0548**	-0.0548**	-0.0548**	-0.115***	-0.0378**
	(0.0221)	(0.0227)	(0.0227)	(0.0177)	(0.0100)
N	650	650	650	582	194
Control group mean	0.56	0.56	0.56	0.54	0.48
Year FE	no	yes	yes	yes	yes
Cohort FE	no	yes	yes	yes	yes
East German + gender	no	no	no	yes	yes
Age groups	29-59	29-59	29-59	32-58	32-58
Cohorts	1954-1966	1954-1966	1954-1966	1954-1966	1959-1962

Source: German Pension Insurance, administrative data on public DI inflows, 1995-2019. Each column in each panel is from one DD model as in equation 1. Panel A also control for East Germany and gender, and Panels B and C control for D_c, T_t but all those coefficients are omitted for readability. See main text for more details.

Table A2: Descriptive Statistic, SOEP Data, 1995-2016

	Mean	SD	Min	Max	N
Panel A. Outcomes					
Public DI I	0.0331	0.1790	0	1	163574
Public DI II	0.0289	0.1676	0	1	163574
Severe health limitations	0.01842	0.134464	0	1	163574
Non employed	0.1865	0.3895	0	1	163574
Full-time employed	0.5951	0.4909	0	1	163574
Individual total income (equivalized)	28,574	30,981	0	2,580,000	163574
Subjective well-being	6.9350	1.7781	0	10	163574
Panel B. Socio-demographics					
Age	44.5985	7.7230	25	59	163574
Female	0.5223	0.4995	0	1	163574
Married	0.7098	0.4539	0	1	163574
Single	0.1289	0.3351	0	1	163574
Children in household	0.9130	1.0672	0	10	163574
Adults in household	0.3596	0.6707	0	7	163574
Household size	1.2726	1.1667	0	12	163574
Dropout	0.0229	0.1496	0	1	163574
Schooling 9 yrs	0.2556	0.4362	0	1	163574
Schooling 10 yrs	0.3595	0.4798	0	1	163574
Schooling 13 yrs	0.2045	0.4033	0	1	163574
Civil servant	0.0594	0.2363	0	1	163574
Self-employed	0.0965	0.2952	0	1	163574
White collar	0.4230	0.4940	0	1	163574
Public Sector	0.2085	0.4063	0	1	163574
Part-time employed	0.2148	0.4107	0	1	163574
In job training	0.0024	0.0491	0	1	163574

Source: SOEP v.33 – 95% sample. Years 1995 to 2016. Only respondents below the age of 60 and birth cohorts 1950 to 1970 are included. See Goebel et al. (2019) for more details about the SOEP.

Table A3: Descriptive Statistic, SAVE Data, 2001-2010

	Mean	SD	Min	Max	N
Panel A. Key variables					
Private ODI	0.3239	0.4680	0	1	12822
Expects Retirement Pre-60	0.02597	0.1591	0	1	12822
Panel B. Socio-demographics					
Age	41.01	10.62	20	59	12822
Female	0.4981	0.5000	0	1	12822
Married	0.6490	0.4773	0	1	12822
Single	0.1926	0.3943	0	1	12822
Children in household	0.8262	1.0383	0	8	12822
Household size	2.5944	1.2643	1	13	12822
Schooling degree 13 yrs	0.4122	0.4922	0	1	12822
Master degree	0.2738	0.4459	0	1	12822
College degree	0.6076	0.4883	0	1	12822
Full-time	0.4786	0.4996	0	1	12822
Part-time	0.1267	0.3326	0	1	12822
Blue collar	0.1756	0.3805	0	1	12822
White collar	0.3343	0.4718	0	1	12822
Self employed	0.0790	0.2698	0	1	12822
Household net income (in 000s)	2.4875	2.4465	0	120	12822
Panel C. Subjective and Objective Health					
Health satisfaction 0-4/10	6.6458	2.4761	0	10	12822
Concerns about own health	0.2011	0.4008	0	1	12822
Smoker	0.3436	0.4749	0	1	12822
SAH	2.4166	0.8377	1	5	9580
Serious Health Issues	0.4564	0.4981	0	1	9580
Heart disease diagnosed	0.0707	0.2563	0	1	9580
Stroke	0.01831	0.1341	0	1	9580
Chronic Lung Disease	0.05481	0.2276	0	1	9580
Cancer	0.0409	0.1982	0	1	9580
High Blood Pressure	0.2292	0.4203	0	1	9580
High Cholesterol	0.13921	0.34618	0	1	9580
# doctor visits	0.6018	0.8131	0	9	8029
# days hospital	0.1926	0.8813	0	27	8029
Normalized health risk score	0.1515	0.1212	0	1	8029
Panel D. Expectations and attitudes					
Subj. life expectancy low	0.2033	0.4025	0	1	8029
Subj. life expectancy high	0.1208	0.3259	0	1	8029
Savings 4 Unexpected	0.7139	0.4520	0	1	8029
Savings 4 OldAge Important	0.7426	0.4373	0	1	8029
No savings possible	0.2034	0.4025	0	1	8029
No savings, enjoy life	0.0242	0.1536	0	1	8029
Higher income expected	2.1876	3.0344	0	10	12822
Inheritance expected	0.8179	2.0289	0	10	12822

Source: SAVE data 2001-2010. Only respondents below the age of 60 and birth cohorts 1950 to 1970 are included. See Coppola and Lamla (2013) for more details about SAVE.

Table A4: Effect of 2001 Reform on Public DI Using Representative SOEP Data

Panel A	*Public DI I* (1)	*Public DI II* (2)	Non-Married (3)	Single Households (4)
Conventional	-0.012***	-0.014***	-0.005	-0.016**
	(0.0038)	(0.0037)	(0.0086)	(0.0077)
Bias-corrected	-0.016***	-0.022***	-0.035***	-0.022***
	(0.0038)	(0.0037)	(0.0086)	(0.0077)
Robust	-0.016***	-0.022***	-0.035***	-0.022*
	(0.0061)	(0.0058)	(0.0134)	(0.0121)
Year FE	yes	yes	yes	yes
State FE	yes	yes	yes	yes
Age & Gender	yes	yes	yes	yes
Panel B.				
Conventional	-0.012***	-0.014***	-0.006	-0.014*
	(0.0038)	(0.0036)	(0.0085)	(0.0077)
Bias-corrected	-0.015***	-0.021***	-0.037***	-0.018**
	(0.0038)	(0.0036)	(0.0085)	(0.0077)
Robust	-0.015**	-0.021***	-0.037***	-0.018
	(0.0060)	(0.0058)	(0.0133)	(0.0120)
Year FE	yes	yes	yes	yes
State FE	yes	yes	yes	yes
Age & Gender	yes	yes	yes	yes
Socio-demographics	yes	yes	yes	yes
Education & labor controls	yes	yes	yes	yes
N	120,211	120,211	34,958	41,434

Source: SOEP v.33 – 95% sample. Years 2001 to 2016. Only respondents below the age of 60 and birth cohorts 1950 to 1970 are included. See Goebel et al. (2019) for more details about the SOEP. The tables shows the point estimates using local polynomial regressions similar to equation (2) (Calonico et al. 2014, 2017, 2018, 2019) using a bandwidth of ten, a univariate kernel, and a quadratic polynomial. Column (2) shows results for an alternative *PublicDI II* measure. Column (3) selects on non-married respondents and column (4) selects on single households. Other robustness checks show results for the pre-reform period (Figure A5), vary the bandwidth (Figure A6), study the smoothness of covariates (Figure A7), carry out density plots of running variables (Figure A8), and vary polynomials as well as carry out donut RDs (Figure A8).

Table A5: Effect on Private ODI Coverage Using Representative SAVE Data

Panel A	Full Sample (1)	Public Pension (2)	No Children (3)	Single Households (4)
Conventional	-0.045	-0.053	-0.017	0.021
	(0.0348)	(0.0443)	(0.0494)	(0.0691)
Bias-corrected	-0.048	-0.051	0.041	0.029
	(0.0348)	(0.0443)	(0.0494)	(0.0691)
Robust	-0.048	-0.051	0.041	0.029
	(0.0417)	(0.0509)	(0.0572)	(0.0794)
Year FE	yes	yes	yes	yes
State FE	yes	yes	yes	yes
Age & Gender	yes	yes	yes	yes
Panel B.				
Conventional	-0.057	-0.075**	-0.060	-0.034
	(0.0464)	(0.0351)	(0.0506)	(0.0671)
Bias-corrected	-0.059	-0.100***	-0.052	-0.010
	(0.0464)	(0.0351)	(0.0506)	(0.0671)
Robust	-0.059	-0.100**	-0.052	-0.010
	(0.0536)	(0.0421)	(0.0596)	(0.0760)
Year FE	yes	yes	yes	yes
State FE	yes	yes	yes	yes
Age & Gender	yes	yes	yes	yes
Socio-demographics	yes	yes	yes	yes
Education & labor controls	yes	yes	yes	yes
N	11,973	9,526	6,236	2,281

Source: SAVE data 2001-2010. Only respondents below the age of 60 and birth cohorts 1950 to 1970 are included. See Coppola and Lamla (2013) for more details about SAVE. The tables shows the point estimates using local polynomial regressions similar to equation (2) (Calonico et al. 2014, 2017, 2018, 2019) using a bandwidth of ten, a univariate kernel, and a linear polynomial. Column (1) is the default sample, column (2) focuses on those eligible for Public DI, column (3) focuses on the childless, and column (4) on one-person households. Other robustness checks vary the bandwidth (Figure A9, Calonico et al. 2020), study discontinuities in covariates (Figure A10), carry out density plots of the running variable (Figure A11, McCrary 2008), and vary polynomials as well as run donut RDs (Figure A12)

Table A6: Health Shocks as Predictors of Labor Market Outcomes: Treated vs. Nontreated

	Public DI (1)	Not Employed (2)	Total Income (3)	SWB (4)
Severe Health Limitation (t-1)	0.0907***	0.0929***	-4,117***	-0.1765**
	(0.0162)	(0.0183)	(623)	(0.0847)
Treated $\times$	-0.0115	0.0397	125	-0.1463
Severe Health Limitation (t-1)	(0.0203)	(0.0252)	(828)	(0.1112)
Treated (t-1)	-0.0056	-0.2161	-17,365	-1.6655**
	(0.0274)	(0.3367)	(14,193)	(0.6866)
N	45,571	45,571	45,571	45,446
R^2	0.0593	0.0314	0.0469	0.0094
Control group mean	0.56	0.56	0.56	0.54
Year + State FE	yes	yes	yes	yes
Socio-demgraphics	yes	yes	yes	yes
Education	yes	yes	yes	yes

Source: SOEP v.33 – 95% sample. Years 2001 to 2016. Only respondents below the age of 60 and birth cohorts 1950 to 1970 are included. See Goebel et al. (2019) for more details about the SOEP. See Burkhauser and Schroeder (2007) for more details about the creation of the *Severe Health Limitations* variable. The indicator is lagged by one period along with the treated dummy that takes one the value one for respondents born after 1960. The dependent variables are indicated in the column headers; column (3) measures total individual income, including various streams of social insurance benefits such as unemployment benefits, sick and maternity leave benefits and all types of pension benefits. SWB stands for subjective well-being.

Table A7: Mean Health Risk Score by Income Quintiles (SAVE)

	Income Q1	Income Q2	Income Q3	Income Q4	Income Q5
Health Risk SAVE	0.1882	0.1648	0.1436	0.1338	0.1196
Health Risk Model	0.1848	0.1583	0.1428	0.1292	0.1056

Source: Tables shows the average health risk score as in Figure 4 by income quintiles. The first row shows the empirical moments form SAVE and the second row those produced by the model.

Table A8: Private ODI Take-Up Rates by Income and Health Quintiles: Policy Simulations

	Health Risk Quintile				
Income Quintile	**Q1**	**Q2**	**Q3**	**Q4**	**Q5**
Panel A: Baseline					
Q1	0.2302	0.2323	0.1158	0.1289	0.0830
...	0.3938	0.3808	0.2783	0.2145	0.1356
...	0.4628	0.4646	0.3725	0.3139	0.2935
...	0.4495	0.4719	0.3539	0.3112	0.2024
Q5	0.4570	0.5016	0.5379	0.4323	0.2590
Panel B: No Admin Costs					
Q1	0.5144	0.6749	0.8956	0.9947	0.5207
...	0.7343	0.7648	0.8892	0.9951	0.7697
...	0.7202	0.8240	0.8191	0.8592	0.6600
...	1.0000	0.9096	0.8795	0.7727	0.6195
Q5	0.7252	0.8812	1.0000	0.9675	0.6239
Panel C: No Private Information					
Q1	0.3304	0.2918	0.5222	0.5196	0.2336
...	0.3834	0.3158	0.7083	0.6622	0.4224
...	0.3952	0.4039	0.8374	0.5031	0.1420
...	0.6185	0.4236	0.8888	0.6367	0.4566
Q5	0.4607	0.4114	1.0000	0.8210	0.5483
Panel D: No Basic Income					
Q1	0.4221	0.6752	1.0000	1.0000	0.9873
...	0.6084	0.5180	0.9212	0.9951	0.9831
...	0.5995	0.3340	0.2979	0.1817	0.8290
...	0.4367	0.2988	0.1785	0.2544	0.6780
Q5	0.5502	0.1949	0.2460	0.2711	0.6158

Table shows private ODI take-up rates for several policy simulations by Health Risk (columns) and Income Quintiles (Rows). Q1 is the healthiest and poorest quintile, whereas Q5 is the sickest and richest quintile. Panel A shows the baseline scenario and replicates Panel B of Table 3. Panel B shows the scenario without administrative costs, Panel C the scenario with full information and Panel D the scenario without a means-tested basic income cash transfer program.

Appendix B: Benefit Calculation

We illustrate the impacts of the 2001 pension reform by running a simple simulation of pre and post-reform benefits assuming a stylized employment history. As explained in Section 2, public DI is a part of SPI. Therefore, we first explain the main method of calculating statutory retirement benefits. Then we explain how disability benefits are calculated.

The German SPI is based on a point system. The gainfully employed earn pension points (pp_{it}) during their work lives. A pension point equals the ratio of *individual* labor income (I_{it}) to *average* labor income ($\bar{I}_t$) in a given year t:

$$\mathrm{pp}_{it} = \frac{I_{it}}{\bar{I}_t} \tag{10}$$

At retirement, the sum of pension points is multiplied by the current "point value" (CPV_t, in €). The value is indexed annually to gross wages and a few other variables. Further, pensions are multiplied by a "pension type factor" (PT_i) which equals one for regular old-age pensions and full WDI pensions. Since 2001, it is 0.5 for partial WDI and ODI benefits. Moreover, there is a fourth factor accounting for actuarial deductions (AD_i) if people retire before the statutory retirement age. Deductions amount to 0.3% per month before reaching the statutory retirement age. The pension, P_{it}, is then calculated as:

$$P_{it} = \sum pp_{it} \times CPV_t \times AD_i \times PT_i \tag{11}$$

DI Benefits. They are calculated like regular old-age pensions. However, as work disability implies leaving the labor market prior to the statutory retirement age, pensions based on prior contributions would be relatively low. Hence disability benefits assume a "reference age." For the period between entry of work disability and this reference age, individuals' *average* pension points are applied. Before 2001, the reference age was 55 and the years until age 60 were valued with $1/3 \times$ average pension points. That is, a person who entered DI at age 40 would get an additional 15 + 5/3 years of her average pension points. Before 2001, there were no actuarial deductions for WDI or ODI ($AD_i = 1$). The factor PT_i was 0.66 for ODI, and 1 for full WDI benefits. Starting 2001, PT_i has been 0.5 for partial WDI and grandfathered ODI, and remained 1 for full WDI benefits.

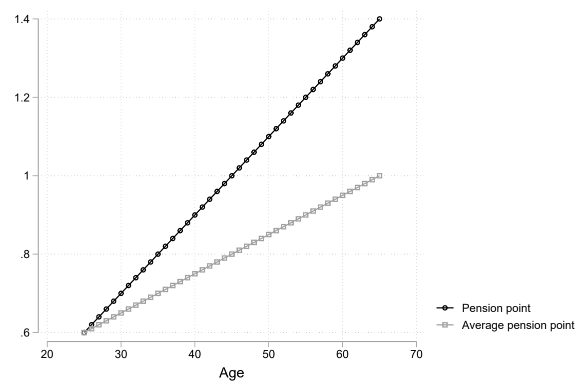
The reform in 2001 also increased the reference age to 60, but introduced actuarial deductions for retirement before age 63. These deductions are capped at 36 months or 10.8% ($AD_i = 0.892$).

As the large majority of disability inflows occur before age 60, the share of DI recipients with maximum deductions of 10.8% exceeds 90%.

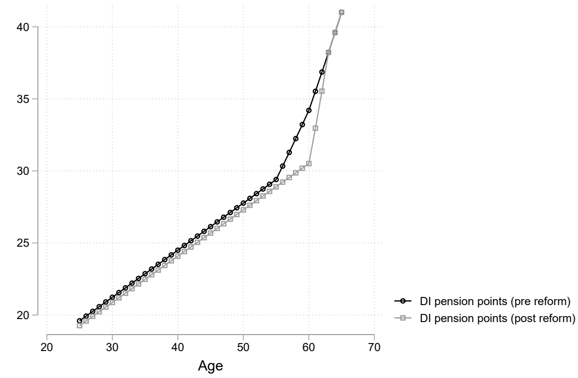
Simulation. Next, we simulate the 2001 reform effect *for benefits and those who were grandfathered in* for a stylized individual. We assume an increasing relative wage position that approximately equals 1 over the lifecycle. The individual starts working at age 25 and earns 60% of the average wage ($pp_{it} = 0.6$). The wage position then increases linearly to 1.4 until age 65 (Figure B1a).

Figure B1: Pension Points by Age and Pre- vs. Post-Reform

(a) Average Pension Points by Age



(b) Total Pension Points by Age, Pre and Post-Reform



Source: own illustration. Note that the post-reform benefits apply either to the grandfathered cohorts who can still claim ODI benefits or the newly introduced partial DI benefits for people who are able to work more than 3 but less than 6 hours a day in any job.

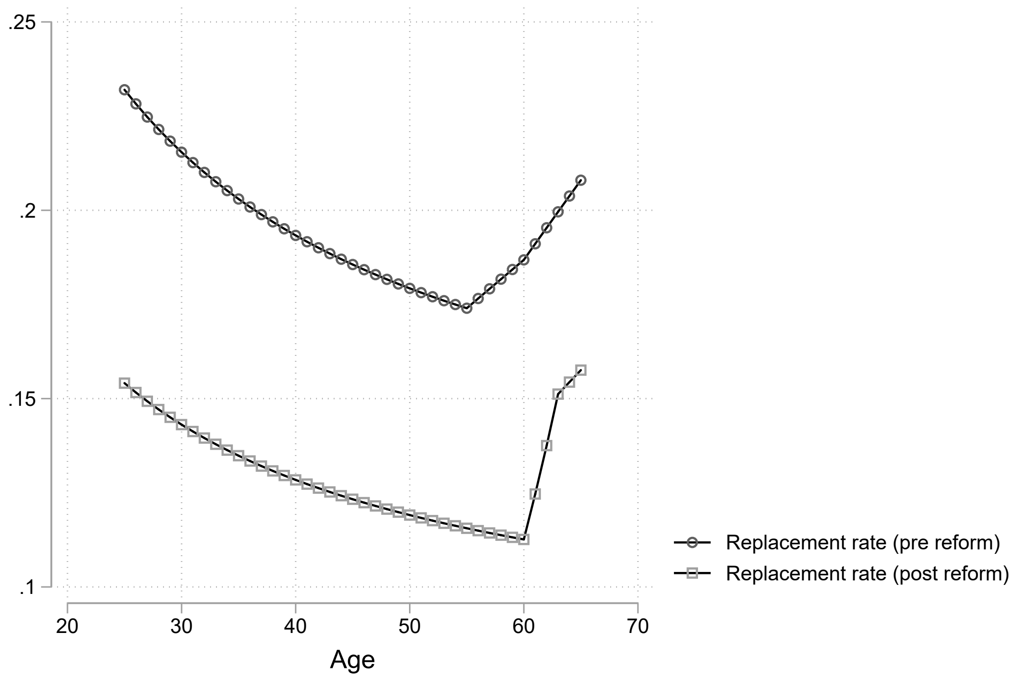
The introduction of actuarial deductions and the increase of the reference age to 60, approximately cancel each other out for most ages. Figure B1b shows that the sum of pension points is slightly lower in the post-reform period. The largest difference applies between ages 56 to 61.[30]

In a next step, we calculate replacement rates by age assuming a single individual without other household income. To calculate the replacement rate, we divide disability benefits by labor income. Figure B2 shows ODI replacement rates in the pre and post-reform periods. Before 2001, the replacement rate was highest at 0.23 at age 25 and then decreased linearly to 0.17 up to the reference age of 55, after which it sharply increased again. After 2001, the general pattern did not change but we observe a downward level shift with a lower replacement rate of between 0.11 and 0.16. Note that these benefit reductions solely applied *for the grandfathered cohorts*. (And for partial WDI, that is, people who are able to work more than 3 but less than 6 hours a day in any

[30] As mentioned, PT_i decreased from 0.66 to 0.5. As a result, benefits—for partial WDI and for the grandfathered cohorts who are still eligible for ODI—are lower as well. The notch cohorts are ineligible for ODI post-reform.

job.) At age 46, the mean age of DI entries, the stylized replacement rate is at 0.18 (pre-reform) and 0.12 (post-reform).

Figure B2: Replacement rate (pre and post-reform)



Source: own illustration. Note that the post-reform benefits apply either to the grandfathered cohorts who can still claim ODI benefits or the newly introduced partial DI benefits for people who are able to work more than 3 but less than 6 hours a day in any job.

Appendix C: Optimal ODI Contracts

This section summarizes optimal insurance contracts in the standard model with private information, when adding administrative costs, and when allowing for a (means-tested) consumption floor. We rely heavily on and refer the interested reader to Braun et al. (2019), especially the proofs therein. For reasons of tractability, assume a single monopolistic insurer and a single risk group that includes a continuum of risk-averse individuals who know that they are either good risk and at the bottom of the disability risk distribution, θ^b or bad risk and at the top θ^t.

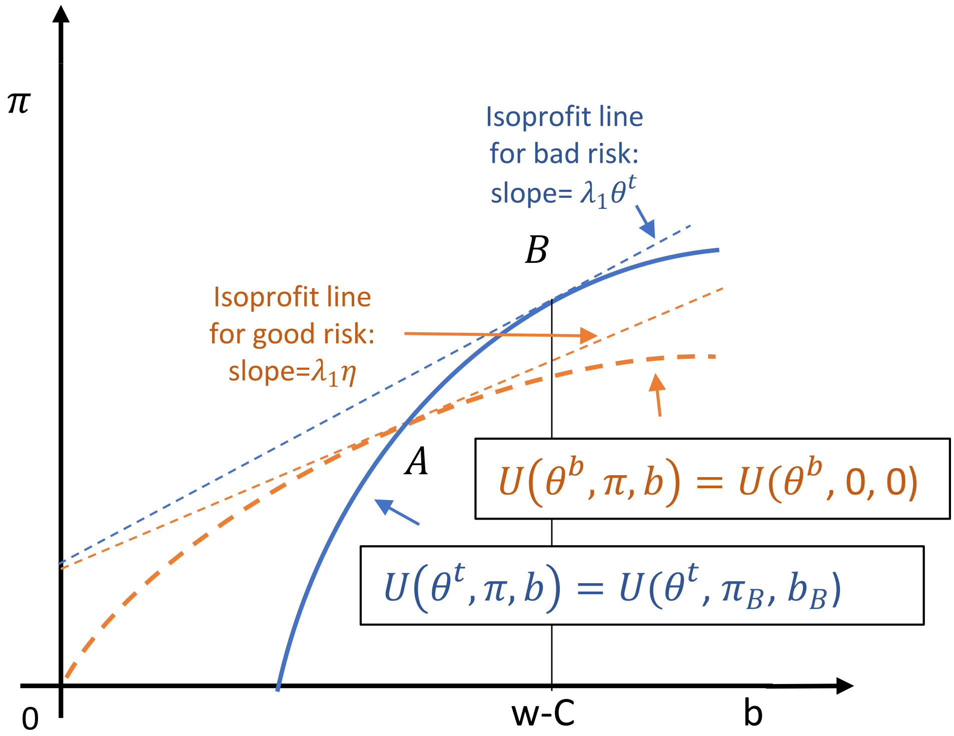
C1 Standard Case: Just Private Information

The core of the standard case goes back to Rothschild and Stiglitz (1976) and Stiglitz (1977). The insurer maximizes profits (see equation (7)), given the participation and incentive compatibility constraints. Figure C1 illustrates optimal contracts under the standard case. The x-axis shows the insured benefit b and the costs of an occupational disability, $w - C$, where w represents the wage in the trained occupation and C is the consumption floor. The y-axis shows the premium Π which increases in coverage levels b.

The flatter indifference curve represents the good risks and the steeper indifference curve represents the bad risks. The slopes indicate the willingness to pay for a marginal increase in benefits. As seen, the bad risks have a higher marginal willingness to pay. The dashed curve that intersects with (0,0) represents the participation constraint when it binds. The participation constraint—indicating that good and bad risks prefer the contracts designed for them over no insurance—binds in the standard case for the good risks. The incentive compatibility constraint—indicating that good and bad risks prefer the contracts designed for them over the other contract—binds in the standard case for the bad risks; the bad risks' indifference curve intersects with the good risks indifference curve. Along the indifference curves, we observe combinations of possible insurance contracts (Π, b) that produce the same utility for individuals, given the participation and incentive compatibility constraints (which are both binding in the standard case).

Consequently, we obtain the optimal contract for the good types where the flatter isoprofit curve of the insurer touches the indifference curve of the good risks at point A. Compared to the optimal contract for the bad risks at B, both the benefits and premium are lower; the contract solely provides partial insurance, whereas the optimal contract for the bad risks in B provides full insurance with $w_o - w_l = b$. We obtain a separating equilibrium.

As discussed, the standard case cannot produce coverage denials by insurers. Only the good



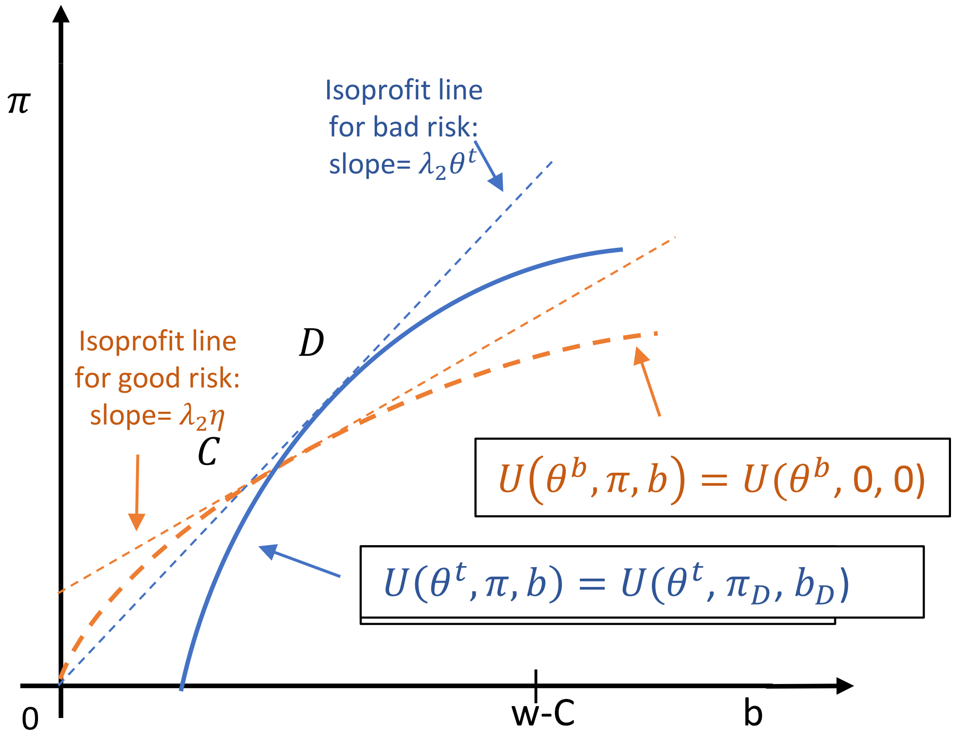
Source:: The dashed indifference curve shows optimal contracts for good risks at the bottom of the work disability distribution θ^b trading off premia (Π) on the y-axis and coverage levels (b) on the x-axis. The solid indifference curve shows optimal contracts for bad risks at the top of the work disability distribution θ^t. The flatter dotted linear line is the insurer's isoprofit curve for the good risks, and the steeper dotted line is the isoprofit curve for the bad risks.

risks can be voluntarily uninsured with (0,0) and produce an ODI take-up that is not 100%. In other words, insurers always offer policies. Such as scenario can happen when the share of the population with low occupational disability risk, ρ, is small, but the dispersion of the true disability risk θ^i—that is unobserved by the insurer—large. In this case, the good types are offered a profitable contract by the insurer, but they prefer to remain uninsured.

C2 Extended Case I: Private Information and Administrative Costs

Chade and Schlee (2020) show theoretically that including administrative costs can produce coverage denials by insurers, as observed in reality. Braun et al. (2019) build on this insight and integrate administrative costs into their model. They show that coverage denials can produce four different scenarios: (i) separating equilibria, (ii) pooling equilibria, (iii) no insurance for anyone, and (iv) and, in practice, a rather unlikely case where only the bad risks are insured.

Once variable administrative costs are introduced, optimal contracts for both good and bad risks never provide full insurance. Further, it could be that all members of a risk group are denied

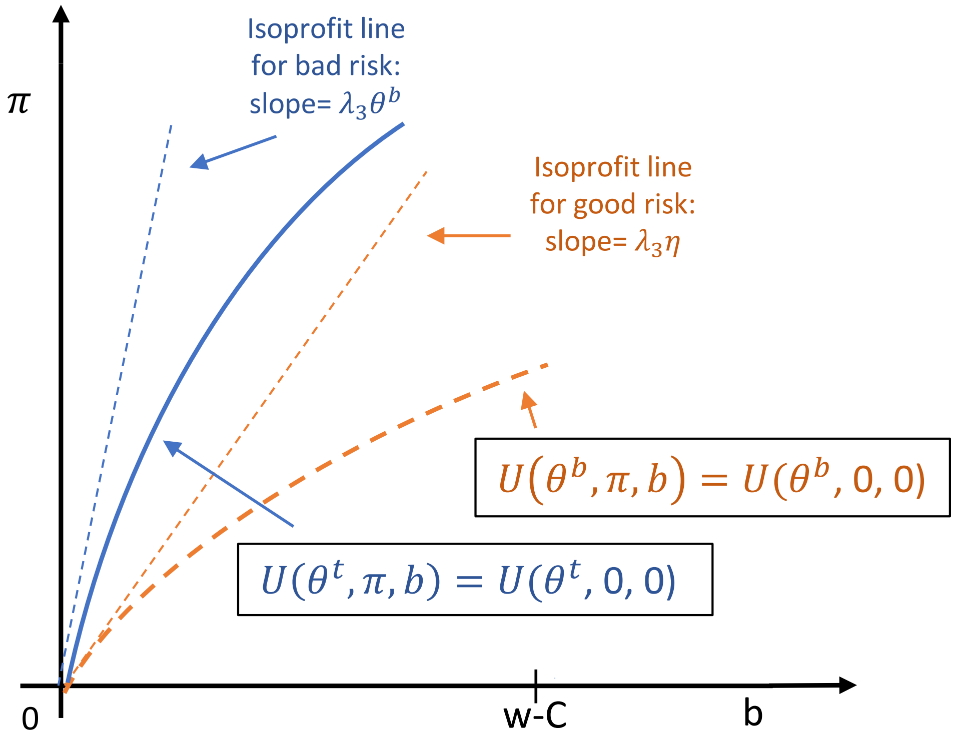


Source:: The dashed indifference curve shows optimal contracts for good risks at the bottom of the work disability distribution θ^b trading off premia (Π) on the y-axis and coverage levels (b) on the x-axis. The solid indifference curve shows optimal contracts for bad risks at the top of the work disability distribution θ^t. The flatter dotted linear line is the insurer's isoprofit curve for the good risks, and the steeper dotted line is the isoprofit curve for the bad risks.

coverage. These are the two relevant cases in practice. As seen in Figure C2, administrative costs lead to steeper isoprofit curves for insurers. This implies that, in a separating equilibrium, the insurer offers policies with lower benefits and premiums. Hence, in Figure C2, optimal contracts for both groups provide less coverage, but also lower premiums (points C and D).

An alternative case would be a pooling equilibrium (not shown), when administrative costs are even higher and where both types are offered the same contract—under the assumption that marginal variable administrative costs are higher for the bad risks. This pooling contract offers even lower coverage, premiums and profits ('skinny plans').

Under certain conditions, when administrative costs are very high, Figure C3 shows a scenario where the entire risk group gets denied coverage. This is because there exists no profitable contract with positive coverage that the insurer can offer. The result is a pooling contract with (0,0) and nobody has insurance. Please see Chade and Schlee (2020) and Braun et al. (2019) for more details and a formal proof.



Source:: The dashed indifference curve shows optimal contracts for good risks at the bottom of the work disability distribution θ^b trading off premia (Π) on the y-axis and coverage levels (b) on the x-axis. The solid indifference curve shows optimal contracts for bad risks at the top of the work disability distribution θ^t. The flatter dotted linear line is the insurer's isoprofit curve for the good risks, and the steeper dotted line is the isoprofit curve for the bad risks.

C3 Extended Case II: Private Information and Social Insurance

Braun et al. (2019) introduce an extension where they include a means-tested public insurer for long-term care costs ('Medicaid') that crowds-out private insurance benefits dollar-by-dollar. This is not the case in Germany where private ODI benefits top-up either the means-tested basic income cash transfer or the basic WDI benefits. This implies that the German private ODI also provides utility when eligible for public benefits, unlike in the US case. Nevertheless, the main underlying mechanisms are the same in the German ODI case: the presence of a public social insurance can lead to optimal contracts with partial coverage. Further, they can lead to the denial of coverage.

Social insurance as a safety net generally increases individuals' utility in the case of no private insurance and thus reduces demand for private insurance; and also profits of private insurers. It increases the individual's outside option and thus the insurer lowers the premium (to satisfy the participation constraint). However, if the consumption floor is large enough, the insurer is unable offer contracts that are still profitable (and provide a sufficiently high utility for individuals). As a result, the insurer denies coverage, see Braun et al. (2019) for details. This case becomes relevant in Germany where the consumption floor is relatively high, especially compared to the initial endowment and occupational disability costs. In this context, uncertainty about future income shocks that may (or may not) result in eligibility for the means-tested basic income affecting demand for private ODI insurance. As explained, we use the representative SOEP to model the income shock distribution over the lifecycle and set the bounds for τ empirically (see Figure 8 and Table 2). As with administrative costs, whether an insurer denies coverage to entire risk groups also depends on the dispersion of private information and the population share of the good risks ρ.

In conclusion, the cutomized general equilibrium model includes multiple risk groups that carry observable h, w, o whereas θ^i is private information. An ODI take-up rate of less than 100% is produced via two different channels. First, insurers deny coverage to entire groups. Second, some individuals are offered a profitable optimal policy but those individuals prefer to self-insure.