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FUNDAMENTALLY REFORMING THE DI SYSTEM: EVIDENCE FROM GERMAN NOTCH COHORTS

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Fundamentally Reforming the DI System: Evidence from German Notch Cohorts
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ABSTRACT

This paper studies how the private DI market and consumers responded to a reform that abolished public occupational disability insurance (ODI) for German cohorts born after 1960. The first part shows a causal reduction in overall DI inflows by more than 30% in the long-run. The second part studies the private individual risk-rated ODI market. Representative data do not show substantial increases in take-up. A general equilibrium model featuring the social safety net, asymmetric information and administrative costs can rationalize these weak private-public interactions as well as stylized facts such as strong private ODI take-up gradients by income and health. It also simulates policies that could have increased take-up further in the course of the reform and assesses their welfare effects. However, although welfare improving, none of the feasible reforms would have increased take-up to more than 60%.

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1 Introduction

How to design social insurance systems has been at the core of economic research (Chetty and Finkelstein, 2013; Cabral and Cullen, 2019; Autor et al., 2019). Three integral strands exist in all OECD (2010) countries: unemployment insurance (Lalive et al., 2015; Hendren, 2017), Workers' Compensation (Powell and Seabury, 2018) and public disability insurance (Kostøl and Mogstad, 2014). What's more, their design and structure are similar across countries. Thus, experiences from one OECD country might hold important lessons for other countries (Besharov and Call, 2022).

A rich literature in economics has studied the implications of DI enrollment for labor supply, earnings, beneficiaries' health and well-being, multi-generational welfare cultures as well as household income, consumption and poverty (Dahl et al., 2014; Deshpande, 2016; Autor et al., 2019; Ruh and Staubli, 2019; Gelber et al., 2023). While reforms that introduced employer experience-rated premiums and gatekeeper protocols to the Dutch public DI system have received much attention (cf Koning and Lindeboom, 2015), there is a lot less evidence on the fundamental German reform of 2001 which cut public DI generosity to move to a mixed public-private insurance system. At the same time, several U.S. reform proposals explicitly point toward the fundamental DI reforms in Germany the Netherlands, arguing that the private market could compensate for reduced public DI generosity (Autor and Duggan, 2010; Fremstad and Vallas, 2013; Winship, 2015; Burkhauser and Daly, 2022). Whether and how private markets can compensate for reductions in public DI generosity is, however, an unresolved economic question.

Our paper examines this fundamental question of how private DI insurance markets and take-up responds to reductions in public DI generosity. To do so, we leverage a German reform that entirely cut public "Occupational Disability Insurance" (ODI) for cohorts born after 1960, effective 2001. ODI insures the lifecycle risk to become work disabled in the *previous occupation* (or a comparable occupation in terms of income and social standing). Public ODI was introduced when the German welfare state was expanded to provide "social status protection" especially for higher income classes. The second strand of German public DI remained intact and is "Work Disability Insurance (WDI)". WDI insures the risk to become work disabled in *any* job.

Using administrative and survey data along with reduced-form approaches, first, we show that the 2001 reform significantly reduced the overall inflow of new DI recipients by more than 30% in the long-run. This first part illustrates that the reform had actual bite among the treated cohorts. The second part studies private market responses. Using representative data and Regression Discontinuity (RD) approaches with birth year as running variable, we do not find much evidence that cohorts whose public ODI was cut purchased private ODI policies at significantly higher rates. Overall, most average estimates let us exclude with 95% statistical certainty that take-up increased by more than 10ppt from a baseline of 32%.

To better understand the underlying economic primitives of the low take-up response, we tailor the general equilibrium model (GEM) by Braun et al. (2019) to the German market regulation. The model features an optimal contracting framework where coverage levels are endogenous and insurers offer two types of policies to high and low risk types, where the "type" is unobserved by the insurer, while pricing is based on observable risk cells and coverage denial common. We model WDI and ODI interactions and leverage three main driving forces: the social safety net, asymmetric information and administrative costs. In addition to rationalizing the weak public-private interaction, the GEM replicates various stylized facts of the private ODI market: (1) Relatively low take-up rates given the high lifecycle risk of occupational work disability. (2) Strong income and health gradients in take-up, and (3) an inversely related, higher risk of work disability among low-income groups and those with high health risks. Finally, we simulate a range of policies that could have increased take-up further had they been implemented along with the reform. Increasing WDI benefits by 10% (while cutting ODI) or measures to eliminate administrative costs dominate from a welfare perspective and would likely have bipartisan support.

Ours is one of the first papers to study interaction effects between public disability and individual private disability markets. Although a rich economics literature studies interaction effects between public and private *health* insurance markets (Cutler and Gruber, 1996; Clemens, 2015), the DI literature on whether and how private markets could substitute for reductions in public DI is much thinner. Institutional differences could be one reason. Although individual private DI exists in Canada and the United States (Autor et al., 2014), the market for group policies, provided by employers, is much bigger but also includes *short-term* DI policies (also called "medical leave" or "Temporary Disability insurance, TDI" in the U.S.), also see Pichler and Ziebarth (2024). Stepner (2021) finds that such employer-provided *private short-term* DI provision has positive spillover effects on *public long-term* DI caseloads. By contrast, Liu et al. (2024) suggest a negative link for the U.S. Outside the U.S., short-term DI—sickness spells of several months while still employed—is typically publicly covered via *long-term sickness insurance* (Ziebarth, 2013; Markussen et al., 2018). As notable exceptions on US group insurance, Autor et al. (2014) study private DI determinants and Cabral and Cullen (2019) use group pricing variation in private supplemental DI to assess the value of public DI. They find that employees value public DI more than twice its costs, in addition to the fact that public DI also insures non-health risks (Deshpande and Li, 2019).

In the broadest sense, the paper contributes to the international literature on disability insurance. For example, Dahl et al. (2014) use the Norwegian case to study financial incentives of DI recipients to return to work—a common cross-country theme in times of rising recipiency rates (cf. Hoynes and Moffitt, 1999; Koning et al., 2022). Dahl et al. (2014) study a return-to-work reform which allowed DI recipients to keep a larger share of their earnings. The program was effective, increased labor supply, reduced program costs, and demonstrated substantial work capacity among DI recipients, in line with Bound (1989) and Maestas et al. (2013). Using a sufficient statistics welfare framework and two Austrian reforms, Haller et al. (2024) show that tighter DI eligibility rules outperform benefit cuts on various dimensions such as fiscal savings. And using quasi-random case worker assignment, international studies find that employment rates among marginally rejected applicants are 10 to 30 percentage points higher compared to marginally accepted applicants (Chen and Van der Klaauw, 2008; von Wachter et al., 2011; French and Song, 2014). In addition to labor supply effects, the DI literature studied how barriers to applying and health screening determine the inflow of DI cases (Autor and Duggan, 2003; De Jong et al., 2011; Deshpande and Li, 2019). Benefit substitution has also been shown to exist in various countries (Borghans et al., 2014; Lalive et al., 2015; Autor et al., 2019; Leung and O'Leary, 2020; Ahammer et al., 2023).

In the broadest sense, the paper contributes to studying the German private DI market. While some papers have described market regulation (Soika, 2018; McVicar et al., 2022), Seitz (2021) estimates a dynamic life-cycle model to conclude that, with a coexisting private market, the welfare-maximizing public DI program would be less generous than in a world without private markets. Börsch-Supan et al. (2022) find that the reform has not systematically improved target quality. Also evaluating the reform and using administrative data, Hanel (2012) does not find that lower benefit levels affected DI inflows. Moreover, in a concurrent working paper, using data from a single private insurer, Seibold et al. (2022) find modest increases in private ODI sales in response to the 2001 reform, advantageous selection, and a low "observed willingness-to-pay of many individuals." In conjunction with a sufficient statistics approach, they argue that distributional concerns and potentially biased perceptions by workers would imply that the 2001 reform was welfare-*decreasing*. By contrast, to study the private market, this paper uses representative survey data and a general equilibrium model(elling) approach in a classical

(Rothschild and Stiglitz, 1976) (RS) optimal contracting framework in a long-run perspective to explain low take-up and stylized take-up pattern more generally .

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 has Statutory Pension Insurance (SPI) (Eibich, 2015; Geyer, 2021) which contains the public DI program (more details below), in addition to a universal means-tested basic income cash transfer program. This means-tested social safety net program provides a guaranteed minimum income floor containingfor a single, monthly cash benefits of €563 Euro (in 2024) in addition to costs for accomodation and utilities.[1] Later in the paper, we will analyze the role of this means-tested basic income program for the low private ODI take-up rate among low-income individuals.

These social insurance programs are funded through a mix of contribution rates for UI, WC, HI, LTC and SPI, employer mandates for paid sick leave, and general taxes for the means-tested basic income program. 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 DI program is part of SPI. It provides benefits for both partially and fully 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 (€7.550 per month in 2024).[2]

[1]For those who are able to work, it is called Unemployment Insurance II (*Arbeitslosengeld II*). 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). 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 treatment and control groups of the 2001 reform (see Section 2.3), but generally cut the generosity of these alternative social insurance strands.

[2]The contribution ceiling is lower in East Germany at €7.450.

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, 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 employees who had paid pension 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. Further, a reform in 1996 introduced caps on the allowed labor market earnings of DI recipients. See Börsch-Supan and Jürges (2012) for a more detailed discussion.

After the 2001 reform (see next subsection), a 2004 reform mandated employers to provide *Workplace Reintegration Management* ("Betriebliches Eingliederungsmanagement", §84 SGB IX). The idea is to overcome temporary disability and to prevent future deterioration 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.

2.3 The Fundamental Public Disability Insurance Reform of 2001

The German public DI system consists of two schemes: (a) work disability insurance (WDI), and (b) occupational disability insurance (ODI). The 2001 reform abolished public ODI for cohorts born after 1960 (our treatment group). Cohorts born before 1961 were grandfathered in (our control group) and were eligible for public ODI before *and* after the 2001 reform.[3] Figure 1 provides an illustration on the main principles after 2001.

The two-tiered system can also be thought of as a combination of a basic (WDI) and supplemental (ODI) scheme. It is important to note that the insurance value of ODI is higher for higher income groups. This higher value for better paying jobs is a function of the eligibility criteria. WDI provides insurance for *general work disability* when a poor health status prevents

[3] 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*.

employees from working *in any job* in the labor market. ODI, by contrast, provides insurance for *occupational* work disability.[4] Importantly, "occupationally disabled" are those who (Deutsche Rentenversicherung, 2023a):

> "[...] due to health reasons, are unable to work in either their trained or a comparable occupation in terms of the education and skills required."

Note that German public ODI was never intended to provide the entire means for a living. Instead it was supposed to provide benefits to compensate for a *partial* loss in work capacity due to which employees either had to switch to a lower paid occupation or from full to part-time work. Obviously, for the most basic and lowest paying jobs in the economy, WDI and ODI converge (Benen, 2023). In fact, the higher insurance value for higher social classes and its inequity implications was a major reason for why ODI was eventually abolished. Note that individual private ODI is also available in other markets such as the U.S. where it is sold as "own-occupation DI" and marketed specifically to higher income professions such as physicians or lawyers (Brian SO Insurance, 2023).

[Insert Figure 1 about here]

Work Eligibility Requirements. The main work requirements to establish eligibility did not change in the course of the 2001 reform, see Table 1, column three. Applicants must have paid pension insurance contributions in the last three out of five years. There has also been a general waiting period of five years throughout the entire time period.

Application & Health Assessment. Details of the application procedure and health assessment are specified in German Social Law and Deutsche Rentenversicherung (2018). Applicants apply at an SPI field office by submitting all relevant documentation such as medical diagnoses and medical records. An independent third-party physician, certified to carry out medical assessments, then reviews the case.[5] Medical reviewers 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 (Deutsche Rentenversicherung, 2023b).

Today, as all grandfathered cohorts have aged out of the possible age to apply, only public WDI exists. The main medical WDI criterion is whether applicants' health limitations prevent

[4] That is, occupational disability "to less than half of that of a physically, intellectually, and mentally healthy person with similar training, knowledge and abilities" (§43, §240 of Social Code Book VI; Viebrok (2018)).

[5] Sometimes these reviewers are state-employed physicians (*Amtsärzte*), and sometimes they are regular specialists practicing in the county of residence of the applicant.

them from working three hours per day in *any* job (Table 1, column three). If applicants' work capacity is less than three hours per day, full WDI is granted (Deutsche Rentenversicherung, 2020). If applicants' work capacity lies between 3 and 6 hours per day, then partial WDI is granted (50% of full benefits). As with ODI, partial WDI intends to compensate for a partial work capacity loss.[6]

Benefits are granted for an initial period of three years and have to be re-certified. After nine years, disability benefits becomes permanent. If work capacity is not expected to improve, permanent DI benefits can be granted earlier, which applies to half of all new cases.

Benefit Calculation. As indicated in Table 1, benefits are calculated as an "early retirement pension" with actuarial reductions. Thus, they 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 (0.3% per month) are applied for everyone receiving benefits before age 63 but are capped at 10.3%.[7] Before 2001, ODI benefits were two thirds of full WDI benefits. [8] After 2001, for the grandfathered cohorts, ODI benefits were 50% of full WDI benefits (Figure 1, column (5)). Appendix B provides detailed pre-post reform simulations of ODI benefits. Post-reform, the simulated replacement rate for a health shock at age 46 is 12% of average pre-DI gross earnings. In practice, in 2000, the average public ODI benefit was €587 per month. The average public WDI benefits was €731 per month (Deutsche Rentenversicherung, 2023b), see Figure 1.

2001 Reform Changes. The crucial and most relevant change in the course of the fundamental 2001 reform was the cut of public ODI for cohorts born after 1960. However, the entire reform package entailed various additional changes, some of which are listed above. Importantly, however, all these additional changes did not affect the birth cohorts differentially and, thus, should not be a threat to the main objective of the reduced-form part—namely, showing that the

[6]However, if part-time jobs are unavailable, partial WDI can be converted to full WDI if the recipient cannot find part-time work within a year. The share of partial converted to full WDI 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.

[7] Several studies documented a high poverty risk among people on WDI (Krause et al., 2013; Märtin et al., 2012, 2014; Geyer, 2021; Becker et al., 2023). 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. Similarly, the age threshold for actuarial deductions has been raised.

[8]However, if reasonable part-time work was unavailable due to the local labor market situation, full benefits could be granted (Viebrok, 2018). 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, case workers would ask the UI office if part-time jobs were available in the region.

reform had actual bite.

For example, the reform also changed how work capacity was medically assessed: from an earnings capacity test[9] to the hour capacity test discussed above. Moreover, as discussed in Appendix B, WDI benefits were slightly reduced for all cohorts, and ODI benefits were reduced for the grandfathered cohorts. Obviously, this reduced benefit level decreased the relative attractiveness of applying for ODI for the grandfatherered cohorts. Thus, the reduced-form estimates on how abolishing public ODI has affected overall DI inflows represent a lower bound estimate.

2.4 Private Disability Insurance in Germany

Basic Principles. 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, 2023), 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.

Coverage Denials. Rating agency data from private insurers covering almost five million ODI policies show 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 (Morgen & Morgen, 2021). 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).

Age at Inception & Claiming. 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:

[9] Pre-reform, the applicant must not be able to earn more than 640 DM (about € 320 in 2001 or $480 today).

mental diseases (32%), musculoskeletal diseases (20%), cancer (18%), and accidents (8%), see Morgen & Morgen (2021).

In our representative data, among those between the age of 20 and 59, about a third of all households with an employee or self-employed person as household head had private ODI coverage (Statista, 2014). In 2015, according to the German Association of Insurers (GDV), the average pension from a private ODI policy was €629 per month (Versicherungsbote, 2020). Our rating agency data yield, on average, an *insured* monthly ODI pension of €1,108 (Morgen & Morgen, 2021) for average monthly premiums of €77. However, in 2014, a high-quality consumer rating report revealed monthly premiums between $50 and $200 for insured monthly benefits of between $750 and $2000 (Ökotest, 2014).

Market Structure. The ten biggest insurers hold more than 60% of the total market share and offer similar but not identical policies (Morgen & Morgen, 2021). 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 hundreds of different actual policies. In an audit study, Ökotest (2014) found very large differences in premiums by occupation and health. Together with healthy profit margins, these facts suggest monopolistic market structures, which we assume later in the model.

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 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 results section, we will further characterize the private German 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) private information as well as (iii) administrative costs 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 2001 reform. That is, we show how it affected the inflow of new public DI recipients and the overall case load. Note that both measures represent *total public DI* cases, the sum of WDI and ODI (as we do not observe ODI and WDI separately in the data). To do so, we use two datasets and two reduced-form identification approaches: (1) Administrative data on the inflow of cases by birth cohort and year in a difference-in-differences (DD) framework. (2) Representative household panel data from the German Socio-Economic Panel Study (SOEP) in a regression-discontinuity (RD) design. Employing the universe of the underlying cohort populations, not just select inflows, we use (2) to validate the findings in (1).

3.1 Impact on Public DI Inflows Using Administrative Data

First, we use administrative data from the SPI to estimate the impact of the 2001 reform on the total public DI inflow. 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 German Federal Statistical Office. Further, we focus on cohorts 1954 and 1966 and ages 29 and 59 in a given calendar year. Then, using data from 1995 to 2019, we compare our treatment group—cohorts who became ineligible for public ODI from January 2001, to our control group—grandfathered cohorts born before 1961. We then estimate the following Difference-in-Differences (DD) model:

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

where y_{ct} denotes the share of new public DI recipients of cohort c in year t; D_c is a treatment dummy; T_t is a post-reform dummy 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 treated cohorts would have developed in the same manner as those of the grandfathered control cohorts. Note that our setting is not prone to possible biases as in staggered DD settings (Goodman-Bacon, 2021).

Results. To illustrate the main findings, Figure 2 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 2 about here]

As seen, whereas the five pre-treatment years show no trending, and the 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.[10] 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. 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 control group. The decline for women is only half as large at 10%. Note that the long-term effect from 2011 onwards is about twice as large (see Figure A2). However, when zooming-in and restricting the bandwidths of the cohorts considered, that is, cohorts born between 1959 and 1962, the effect sizes decrease to -12.5% for males and -7.9% for females—on average over *all* post-reform years 2001-2019. 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.[11]

[10] 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.

[11] In addition, recall that many ODI pensions were converted to full WDI pensions if recipients could not be referred to a "reasonable" job, see Section 2.

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 (SOEP) and an alternative identification approach. The SOEP allows us to observe *representative* samples of each 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 now focus on the main approach and findings.

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

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

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

where y_{it} indicates whether the respondent receives public DI benefits. D_i is one if the respondent belongs to the treated 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, 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). In the main results, we present robust and bias-corrected estimates (Calonico et al., 2018). In the appendix, we vary the bandwidth, use data-driven bandwidth selection (Calonico et al., 2020), and covariates (Calonico et al., 2019).[12] Moreover, our estimates are robust to implementing methods for discrete running variables following Kolesár and Rothe (2018).

The main RD identification assumption implies that no other factor would have affected

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

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*.[13] 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 1 very well.

Results. Figure 3 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. Figure A3 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.[14]

[Insert Figure 3 about here]

Table A3 shows the RD results using local polynomial RD methods for the post-reform period 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 3 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

[13] 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.

[14] 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.

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. Next we ask the question: For the treated cohorts without access to public ODI, how does a health shock materialize after as compared to before the reform, 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 a binary indicator for severe health limitations, 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 year after a severe health shock, (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.

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%, cf. 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 3 and Table A3 (less likely to be on public DI), it is imprecisely estimated. Similarly, the interaction effects suggest (imprecise) increases in non-employment by about 4ppt, and small and insignificant effects for changes in income and well-being.

4 The 2001 Public DI Reform and the Private DI Market

Next, we study interaction effects between the public and private ODI market. Moreover, we provide general insights into the functioning of 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 in Section 5, building on Braun et al. (2019), we use a tailored version of their general equilibrium model (GEM) to explain these stylized empirical pattern.

4.1 Impact of the 2001 Reform on the Private ODI Market

In a first step, we investigate whether the treated cohorts purchased private ODI policies at higher rates, relative to the control 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.

Data. We use representative survey data from SAVE (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 the relevance of 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.[15] Table A4 shows the summary statistics of our main sample. 32% of all households are ODI policyholders, the average age is 41 and 41% hold the highest schooling degree in Germany after 13 school years. To identify the treated cohorts, we directly observe the birth year as a separate variable.

Figure 4 illustrates the main result on private ODI take-up for the full sample; Figure A8 (Appendix) shows robustness checks for alternative samples; clockwise, starting from the upper left: (1) the full sample as in Figure 4, (2) those eligible for public DI, (3) childless households, and (4) one-person households. In all graphs, 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,

[15] We ignore civil servants who were not affected by the DI reform.

overlaid with linear plots for each side of the cut-off.[16]

[Insert Figure 4 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 have private ODI. This is not surprising. 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 1. 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. Using time series data on all sold private ODI policies in Germany, Besharov and Call (2022) show a substantial expansion of the private ODI market long before the 2001 reform, in line with this conjecture. Further, younger applicants are also 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 treated 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 4, 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 to -0.06. Overall, in Table A5, 19 out of the 24 estimates carry negative signs, not the hypothesized positive ones.

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 does not alter the findings (Kolesár and Rothe (2018), detailed results available upon request). However, the robustness checks also illustrate that most point estimates carry relatively large standard errors. Nevertheless, in column (1) of Table A5, we can exclude with 95% statistical

[16] Linear slopes fit the data better than quadratic ones; however, we vary polynomials in robustness checks.

certainty that the treated cohorts took up private ODI insurance at a rate higher than 3 to 5 percentage points (ppt), relative to the baseline of 32% (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 A3, the increase in take-up was at most 5 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 could 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, Section 5 employs a GEM based on Braun et al. (2019) to trace out the underlying driving forces for the low post-reform take-up rates and interaction effects between the public and private DI system. First, however, we now present several stylized facts about private ODI take-up. The GEM below is powerful enough to rationalize the weak private take-up pattern and these stylized facts. To produce the stylized facts, we rely again on the representative SAVE and SOEP surveys.

Health Risk Score. Table A4 shows a detailed list of health measures 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 coverage denials.

[Insert Figure 5 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 5. It is reassuring to see a typical left-skewed health risk distribution with a long right rail as common for health risk distributions (Karlsson et al., 2016, 2024; Atal et al., 2023).

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 occur an occupational disability. The average age when becoming work disabled is 46, and the average ODI contract runs until age 64. To do so, we focus on a sample of SOEP 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 ODI policies. We then cut a SOEP lifecycle sample such that we also observe *the same individuals* at least once between the age of 55 and 60. Following Burkhauser and Schroeder (2007), we elicit a broad and representative measure of the lifecycle risk of occupational disability among German employees.

[Insert Figure 6 about here]

Lifecycle Risk for Severe Health Limitations. Figure 6a plots the lifecycle risk of having a severe health limitation (at least once) against the quintiles of self-reported health satisfaction between 25 and 35. We use health satisfaction as a proxy for health when entering the labor market (Ziebarth and Karlsson, 2010).[17] It also stratifies the risk by the quintiles of household net income. We summarize Figure 6a as follows: First, the lifecycle risk of a severe health limiting shock is large. Second, it remains significant even for the healthiest employees. On average, 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 6b 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 high, above 20% for most health and income groups. For example, averaged over all health quintiles, the lowest income quintile carries a lifecycle risk of 37%, the second lowest of 24%,

[17] 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.

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.

ODI Take-Up by Health and Income. Figure 7 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 in Figure 5 on the x-axis. The downward sloping lines are again stratified by quintiles of net household income.

[Insert Figure 7 about here]

We can summarize: First, 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 get charged health risk add-ons.

Second, Figure 7 shows that, across the entire health risk distribution, the lowest income quintile has take-up rates that are substantially lower than all other income quintiles, between 25% for the healthiest and below 10% for the sickest people. In other words: The poorest 20% of the population have private ODI take-up rates of just 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: First, 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, this lifecycle risk increases with worse health and lower income. Nevertheless, private ODI take-up rates are relatively low and between 10% and 50%, given this high lifecycle risk, even after substantial reductions in public DI generosity. However, importantly and paradoxically, take-up is *inversely correlated* with the lifecycle risk as the sickest and poorest have the lowest take-up rates despite having the highest work disability risk.

[Insert Figure 8 about here]

Private Information. Finally, we exploit the rich SAVE data to provide evidence for private information. Figure 8 plots private ODI take-up rates on the y-axis and the five health risk score quintiles on the x-axis, as before. However, now we plot two lines differentiating by whether SAVE respondents *expect to stop working before age 60* which proxies for expected work disability. As seen, over the entire health distribution, those who expect work disability have substantially higher private ODI take-up rates. While this empirical pattern is no 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 (GEM) by Braun et al. (2019), which is based on Rothschild and Stiglitz (1976), Stiglitz (1977) as well as Chade and Schlee (2020). Braun et al. (2019) develop a variant of Rothschild and Stiglitz (1976), as in Stiglitz (1977), to study in an optimal contracts framework with endogenous coverage the role of three key factors in long-term care insurance take-up in the US: Medicaid (public insurance for low-income individuals), administrative costs as well as private information. 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.

We build on Braun et al. (2019) and the literature above, but customize the GEM to capture the institutional details of the German private ODI market. The model by Braun et al. (2019) is flexible and powerful enough to explain both, the weak public-private market interactions as well as stylized take-up pattern in the German private ODI market. To do so, the GEM leverages three main driving forces: (i) The German means-tested basic income program, (ii) the public basic WDI program, (iii) private information, and (iv) administrative costs.

The tailored GEM captures the main regulatory framework of the German private ODI market, which leads to high denial rates for certain risk groups. Appendix C discusses optimal contracts and market equilibria. Note that the model is a standard GEM; 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 purchase private ODI insurance at an average age of 32. The average age when health shocks lead to occupational work 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 9.

[Insert Figure 9 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):[18] $c_1 = w_1 - s$. Health, wages and occupation are jointly distributed with density $f(h, w_1, o)$.

Period 2: ODI Offers and Purchase Decisions. While insurers observe health, wages and occupation, those are noisy indicators of the true occupational 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 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. 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.

Individuals may or may not purchase ODI that tops-up household income due to several reasons.[19] First, in contrast to the insurer, they know whether their true disability risk is high or

[18] 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., 2023). 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.

[19] The "means-tested basic income program" and "public WDI" set very similar consumption floors and the former

low, but they do not know their actual risk with certainty. They weigh the risk of occupational disability against paying a monthly premium Π to insure this risk.

Second, the *participation constraint* (see below) ensures that each type i prefers, if offered, the specifically customized ODI policy over no insurance.[20] 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 to all residents. Thus, absent uncertainty, 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 also become eligible for the means-tested basic income program if their income drops sharply, which is uncertain in the model.

Period 3: Income and Health Shocks. Period 3 represents individuals' main work lives and stretches from age 35 to retirement. In Germany, the earliest possible age to receive a statutory early retirement pension is 62. Individuals are aware that future wages 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.

A health shock may either lead to work disability (WDI) or occupational disability (ODI).

WDI Costs. In the WDI case, individuals cannot earn any labor income but receive public WDI benefits which are roughly 30% of their previous income, see Section 2 for details on the calculation. Thus individuals' costs of work disability are 70% of their previous wage. Because WDI benefits are calculated like an early retirement pension as a function of previous wages, they do not differ by health risk or occupation. Note that WDI recipients also receive full private ODI benefits which are not crowded out. As WDI implies a larger health shock than ODI, those with a private ODI policy receive benefits in case of full work disability.

ODI Costs. Occupational disability, by contrast, implies that individuals cannot work in their previous occupation anymore, or only part-time. We assume that individuals with an ODI shock earn half their previous wage, see below for more details.

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 €730 in 2005, see Figure 1. In 2022, after various benefit increases over time, see Section 2, it was basically identical to the basic income level and €1,007 (Deutsche Rentenversicherung, 2023b).

[20] 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.

Costs of Disability. Thus, the monetary costs of a health shock that either leads to work or to occupational disability, $l(h_i)$, are:

$$l(h_i) = P(WDI|WDI \vee ODI) * 0.7 * w_2 + P(ODI|WDI \vee ODI) * 0.5 * w_2 \tag{3}$$

WDI vs. ODI Probability. However, the probability that either occupational or work disability is triggered by a health shock differs by occupation and health risk. Specifically, if employees in the most basic (low wage) occupation occur a work limiting health shock, it will always lead to WDI as there is no wedge between a high(er) occupational status and being able to do the most basic work in the economy. By contrast, given a health shock, we assume that employees in the most skilled, highest status, occupations have a risk to occur full work disability of 20% whereas the occupational disability risk would be 80%. We extrapolate the conditional ODI risk for the intermediate categories. Our assumed work vs. occupational disability probabilities by health risk quintiles from category 1 to 5 are $P(WDI|WDI \vee ODI) = [0.2, 0.4, 0.6, 0.8, 1]$. Inserting those probabilities in equation (3), for a work limiting health shock, we obtain the following expected costs as a share of previous wages by health risk quintiles: $l(h_i) = [0.54, 0.58, 0.62, 0.66, 0.7]$.

Consumption Floor. After any work disabling health shock and its resulting wage losses, following reality, the model checks whether the remaining work capacity produces enough market income to earn above or below the means-tested basic income threshold of about €1000. For example, assume a highly qualified white collar worker with a monthly income of €10,000 has a health shock. In expectation, she occurs a 54% wage loss, see above, and her income drops to €4,600. If she holds a private ODI that pays a monthly ODI benefit of €2,000, her income after a health shock would be €6,600, and the average disability costs would be €3,400.[21] However, assume a low-income employee in the second highest risk quintile with a monthly income of €2,000 has a health shock. In expectation, she would lose 66% of her income, as $l(h_i) = [0.54, 0.58, 0.62, 0.66, 0.7]$. Thus she falls below the basic income threshold and receives C.[22] Thus, in case of a work limiting health shock, those without private ODI occur average costs of $\min[w_2 - C; l(h_i) * w_2]$, and those with a policy $\min[w_2 - C + b - \Pi; l(h_i) * w_2 + b - \Pi]$ resulting in consumption c_{ODI}, where

[21] Under WDI, her income drops to €3000 without and to €5000 with private ODI. Under ODI, her income is cut in half to €5000 without private ODI benefits and to €7000 with private ODI benefits.

[22] The average German gross wage was €47,928 in 2019 (Statistisches Bundesamt, 2022), and the average insured benefit for occupational disability was €13,301 p.a in 2019 (Morgen & Morgen, 2021). By our assumption above, a health shock then leads to an average wage loss of 62%, or €18,213. Those who hold private ODI would earn a total income of €31,514.

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

and r is the real interest rate on savings, s. Ψ is a social insurance transfer via the German means-tested basic income program, where $\Psi = max[0, C - ((1-\tau)w_2 + (1+r)s - \Pi + b)]$. Individuals who buy a private ODI policy but do not occur any disability, pay premia Π but incur no wage losses:

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

In reality, the private ODI market provides hundreds (if not thousands) of different policies. Experts recommend to insure an income level of 70% of the gross wage; 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

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

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:

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

where variable insurer costs are λ (e.g. claims processing) and fixed insurer costs are γ (e.g. broker commissions). The incentive compatibility constraint ("I prefer my policy over the policy designed for the other type") is

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

and the participation constraint ("I prefer my policy over no policy") is

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

5.1.3 Parameterization of Model

We follow Braun 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. An important first step is to model the representative health risk score distribution in Figure 5. A beta distribution with $\beta(1.2269; 6.9219)$ approximates this skewed distribution reasonably well. As illustrated in Figures 6 to 8, 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. 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 35. We model it as a log normal distribution and normalize Period 1 (Figure 9) 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.[23] Similarly calculated are the costs 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 9.[24] 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, 2023).

[23] 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)).

[24] (€47,928-€11,868)$\times$16/(€47,928$\times$31.5)

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 is that the menus of optimal insurance policies that insurers offer must be calculated for each of 750 different risk groups. The risk groups consist of combinations of health risk (h), income (w), and occupational groups (o).

We simultaneously set parameters to minimize the distances between the actual data moments and equilibrium model outcomes: the distribution of income uncertainty τ, the work disability risk θ 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 6.[25] 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 powerful enough to replicate the stylized empirical facts of ODI take-up that we observe 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 ODI insurance that provides little insurance value and thus utility, given that they likely qualify for the means-tested basic income program after an income or health shock.

[25] 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).

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 7. 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 7).

[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 produces a private ODI take-up rate of 25.7% instead of 25.9% for the poorest but healthiest quintile; for the healthiest and richest quintile it is 47.9% instead of 45.2%, and for the poorest and sickest quintile it is 11.4% instead of 9.6%.

5.1.5 Simulating the Reform

Next, we simulate the 2001 reform. Unfortunately, no pre-reform SAVE data exist. Consequently, we use the model to simulate the *reverse* reform effect. To do so, we require an assumption about the pre-reform replacement rate for public ODI to specify the pre-reform costs of an occupational disability in percent of former income. Appendix B illustrates 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 then increases linearly to 140% if the individual worked until age 65. In addition, using representative SAVE data, we apply an alternative assessment using information about the individual expected pension replacement rate as well as the current wage. Reassuringly, this methods produces an almost identical average replacement rate.[26]

Figure 10 shows an average simulated pre-reform private ODI take-up rate of 19% and an average post-reform take-up rate of 33%. In our representative SAVE data, 22% of all households who were born prior to 1961 have private ODI, and 42% of all households born after 1960 have a

[26] Specifically, we exploit a SAVE question about the expected statutory pension replacement rate as a share of the last net wage. As we do not know the precise individual work history, exploiting this individual-level information succinctly summarizes individual knowledge and expectations about future pensions. Then, following the institutional details, we assume that the public ODI pension would have been half the expected statutory pension and we also apply a 10% actuarial deduction as the large majority of recipients receive public ODI before age 60, see Section 2 and Appendix B. To calculate hypothetical public ODI replacement rates, we relate the hypothetical public ODI benefit (average: €341) to (a) the current gross wage (average 17%) as well as the (b) current net household income (average 12%). We use the latter as default but results are similar when using the former.

private ODI policy, see also Figure 4. However, these simple averages ignore cohort effects, time trends or age differences in take-up and do not represent causal estimates. Also not that the 95% confidence intervals of our estimated reform effect, including our sets of robustness checks, include the model-based simulated reform effect (Table 1 and Figure A12).[27] When estimating reform effects by the five health risk quintiles, we find imprecise estimates for quintiles two to five, but a significant 22ppt take-up rate among the healthiest quintile of the population. The simulated model take-up rates for health risk group 1 is 10% and for health risk group 2, it is 16%, whereas it is zero for the sickest quintile (detailed results available upon request).

Finally, it it worthwhile to mention that these simulations are not identical to the empirically elicited estimates. In the GEM model, the decision of purchasing is made at young ages when entering the labor market, whereas the empirical estimates pool all individuals below 60. Importantly, the model abstains from moral hazard which would, in reality, reduce expected ODI costs and thus lower the impact of cutting public ODI on private ODI take up.

Overall, we conclude from these multiple sources of evidence that our model predictions map up well with external data sources.

5.1.6 Alternative Policy Simulations

Finally, we use the model to simulate policy counterfactuals; we always combine the actual reform with specific policies and then simulate take-up effects under these hypothetical scenarios. Specifically, we simulate how the reform would have changed the private ODI take up rates—in the general population and by health—had policymakers it combined with the following measures:

1. An increase in WDI benefits by 10%.
2. A reduction of the consumption floor by 10%.
3. No variable administrative costs, e.g. through minimum benefit ratios.
4. No fixed administrative costs, e.g. through banning broker commissions.
5. Elimination of private information, e.g. through genetic testing.

The results are in Figure 10. As seen, the increase in take-up would have been one percentage points larger had policymakers simultaneously increased WDI benefits by 10%. The would have

[27] Using data from a single insurer and difference-in-differences methods, Seibold et al. (2022) estimate a reform effect of 18ppt.

been a reasonable and likely bipartisan policy option. Increasing WDI benefits increases the value of ODI policies, especially for low-income populations as it becomes less likely that the means-tested consumption floor crowds-out private ODI benefits (recall that public and private ODI benefits do not crowd each other out in the German setting).

[Insert Figure 10 about here]

Further, a reduction in the means-tested German safety net by 10% would have increased take-up by 23 percentage points to an overall coverage rate of 42%. Here, again, the intuition is that such a measure would increase the value of private ODI policies for large parts of the population as the consumption floor decreases. An extreme thought experiment would be to entirely abolish the German safety net (which would be unconstitutional). In such a scenario, private ODI take-up rates would have more than tripled to 81% in the population (detailed results available upon request).

The next two bars in Figure 10 illustrate the increase in take-up without fixed or variable administrative costs. Note that either heavily limiting or even banning broker fees is a realistic policy option that is under active debate not just in Germany (Bundesanstalt für Finanzdienstleistungsaufsicht (BaFin), 2018) but is, in fact, under active consideration at the level of the European Union (Reuters, 2023). Lobby groups but also academics sometimes criticize such regulatory measures arguing that unintended cost-shifting or shifting of business activities could be a consequence (Inderst and Ottaviani, 2012; Braegelmann and Schiller, 2024). However, in the U.S., when the Affordable Care Act was introduced, the regulator implemented so called "medical loss ratios" to limit administrative spending by insurers (Born et al., 2023; Statistisches Bundesamt, 2024). In any case, according to our simulations, take-up would have increased to 45% (no variable admin) and 46% (no fixed admin), respectively. The increase through lowering admin costs is bound at 54% for the unrealistic case of zero admin costs (results available upon request).

Another, rather unrealistic upper bound scenario is the last bar in Figure 10. It indicates the increase in take-up had the reform been combined with measures to eliminate information asymmetries. One extremely controversial policy option that would not even achieve an complete information asymmetry elimination could be genetic testing. Take-up would have been 64% in that case.

Further, Table 4 lists how other relevant outcome measures would have evolved, both for the general population and by the two (unobserved by the insurer) risk types. We differentiate between good and bad risk types, that is, θ^b and θ^t. As shown in Table 2, 73% fall into the bottom of the disability risk distribution. Most importantly, while observable risk factors like health, health care use, income and occupation is priceable in this market, the risk type is unknown by the insurer. Thus, in Table 4 , we observe adverse selection with only 3% of good risks purchasing policies, but 45% of bad risks. Further, the share of costs covered is much higher for bad vs. good risks. High types (bad risks) receive relatively favorable terms because insurers design contracts under incentive compatibility constraints, which limit the amount of profit they can extract from them.

Plan Generosity. Interestingly, while the share of risk covered by ODI policies remained basically constant over the reform at 84% for bad risks, for good risks, it increased from 29% to 40% through the reform illustrating that good risks demanded and were offered more generous ODI policies. In addition, the reform reshuffled loads from bad to good risks, that is, bad risks profited through a lower loading factor along with higher generosity plans. The loading factor 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.

Loading. Pre-reform, insurers would load contracts for bad risks substantially (0.66), whereas the loads for good risks were at 0.24, that is, we observe cross-subsidization. The reform essentially flipped the roads by type. Further, policy options to eliminate variable and/or fixed admin costs would further reduce the load for bad risks (and increase it for good risks)—along with a higher take-up to around 80% under each policy for bad risks.

Equity. Figure A13 (Appendix) takes another cut at the simulations by studying take-up by (observable) health risk quintiles. The lack solid line shows take up post-reform; the other lines illustrate how take up would have been evolved if combined with alternative policies, see above. We conclude: First, not surprisingly, take-up never exceeds 60%. Second, none of the combined policies achieves the elimination of the health gradient. Eliminating fixed administrative costs, that is, broker commissions (but also "variable admin costs") comes close to that potentially desirable goal but fails for the sickest health group. In general, basically all alternative policies fail in increasing take-up among the sickest health group. By contrast, reducing the basic income floor would have left the gradient intact but shift take-up upward by between 5 and 12 percentage points.

In conclusion, our simulations illustrate that policymakers could have undertaken reasonable and realistic policy measures to increase private ODI take-up above the very modest reform-induced increases. The most promising approaches would have been an increase in WDI benefits or regulatory limitations of administrative costs, for example, through a bad of broker commissions as envisioned by the European Union Reuters (2023).

When studying how these policy options would have affected different health groups, that is, affected equity, in particular the elimination of variable administrative costs (but also fixed costs) would have reduced the health gradient in take-up substantially, see Table 4 and Figure A13. Recall that administrative costs are substantial in this individual market—fixed costs like broker commissions amount to an estimated 3% of lifetime premium payments over 31.5 years, in addition to 10% 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.

Welfare. Next, we discuss welfare which includes consumer utility, insurer profits and government spending. Table 4 lists insurer profits along with consumers' ex ante utility. Note that this utility includes utility from insurer income transfers (and thus insurer profits) and allows us to assess welfare in a compact manner. Note that, because of CRRA with a risk preference factor of 2, utility is negative and that less negative utility implies an improvement. In total, we display three rows of transfers: (i) DI transfer, (ii) basic income transfer and (iii) total transfer, the sum of (i) and (ii).

As seen, eliminating administrative costs would reduce the basic income transfer by almost the same amount as reducing the consumption floor directly by 10% or increasing WDI benefits by 10%. All these measures would increase take-up and private ODI benefits in the case of work disability, thereby reducing safety net payouts. However, "DI transfer" is a much larger factor compared to the basic income transfer; all policies reduce public DI spending in a similar fashion, substantially, as a result of the reform.

Assuming an average ODI benefit as shown in Figure 1 and a decrease in inflows by a third per cohort per year, the reform reduced expenditures of the German Statutory Pension Insurance by €789 million in the first year, or about €12 billion after 15 years, as the first affected cohorts aged through their work lives. Assuming that the reduced spending would be passed entirely through to SPI contribution rates, it would result in a reduction of one percentage point. This translates to a reduction in employer and employee taxes by €500 per year for an annual income of €50,000.

Despite rising insurer profits through the reform, overall consumer utility (which includes insurer profits) decreased as public ODI eligibility was cut. However, reduced consumer utilitydoes not consider the reduction in public DI transfers which also decreased substantially (Table 4), indicating that the reform was, on net, welfare improving by these metrics.[28] However, welfare could have been higher had policymakers combined the reform either with a 10% increase in WDI benefits or measures to eliminate either fixed or variable administrative costs.

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 safety net benefits by 12%, its asset eligibility thresholds by 50% (Deutscher Bundestag, 2022), WDI benefits in 2014 and 2019 (see footnote 7), and also reduced possibilities to sanction those who receive cash benefits but are unwilling to cooperate with caseworkers, political reforms to lower the consumption floor are likely politically infeasible. In a 2018 representative poll, 53% of Germans found the basic income benefit levels inappropriately low (YouGov, 2018). On the other hand, policymakers could have simultaneously increased WDI benefits that depend on previous earnings. Strengthening this benefit would have increased welfare and would have been a bipartisan and likely voter-popular option.

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—like none of the policy options—would not eliminate the strong health and income gradients in coverage.

The final option, reducing administrative costs, clearly emerges as the most desirable and feasible of all policy alternatives. Our findings show: the potential to increase take-up is substantial and the health gradient in coverage would be substantially flattened when eliminating variable administrative costs. 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.

Nevertheless, while all these additional policy measures would have increased take-up above the current levels, none of the measures would have been well suited to increase take-up

[28] This assumes that all employees and employer contribution rates would have decreased by about €500 per year on average; in other words, it assumes that the reduced public spending would result in lower contribution rates (and not in lower general tax subsidies which currently exceed €100 billion per year).

above 60% at the population level. If policymakers' objective was to achieve coverage rates above 60%, more fundamental policy reforms such as an individual mandate and guaranteed issue would be necessary.

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) for cohorts born after 1960 starting 2001. This paper first studies the first-stage effects on public DI inflows to demonstrate that the reform had "bite." 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 more than 30% in the long-run, we do not find much evidence that the treated cohorts purchased private ODI policies at substantially higher rates than the control cohorts. Across various robustness check specifications, we can exclude increases of more than 10 percentage points at the population level. As these small to at best modest public-private interaction effects are an equilibrium outcome, we tailor a general equilibrium model to the German regulatory ODI framework. 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 take-up pattern in the private ODI market: (1) despite a high lifecycle occupational disability risk, low take-up rates of less than 50% across all health and income groups. We find at best modest interaction effects with public DI. (2) Strong positive take-up gradients in good health and higher income that are (3) inversely related to the negative gradients in occupational disability risk.

Next, our policy simulations suggest that policymakers could have increased take-up more by either (i) streamlining the means-tested basic income program (which would be politically difficult and may have unintended consequences), (ii) increasing WDI benefit levels, or (iii) 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). Finally, policymakers could (C) implement reforms to lower 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 cap or even ban such high commission fees (Reuters, 2023). Alternative policy proposals could impose "benefit loss ratios" akin to the regulation of U.S. health insurers through the Affordable Care Act that imposed medical loss ratios. Targeting administrative costs could not just substantially increase take-up rates across the entire risk distribution but also reduce the health gradient in take-up, even under risk rating. Our simulations predict that the share of costs covered by private ODI policies (as well as take-up through less pricy policies) would increase for both types and loads decrease for bad risks.

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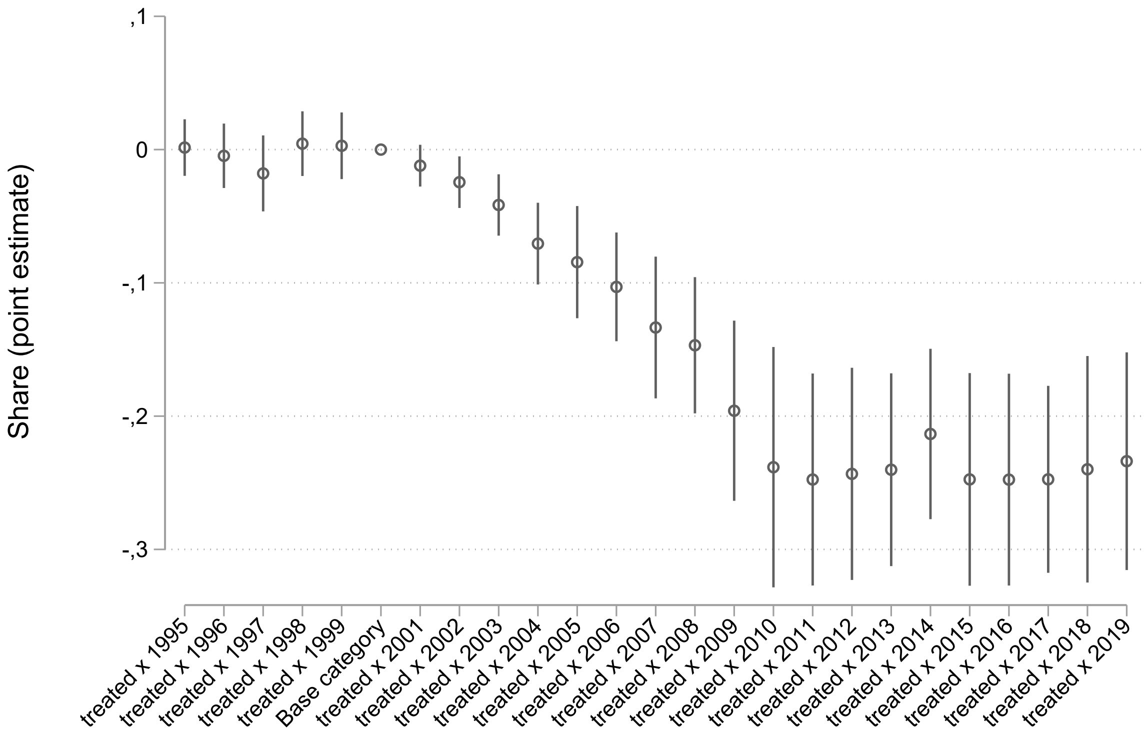
Figure 1: Illustration of WDI and ODI Schemes

Scheme	*Main criterion*	*Work eligibility*	*Health Assessment*	*Benefits*	*Calculation (Appendix B)*	*Notes*
Work DI (WDI)	Work disability in any job	Social contributions paid in last 3/5 years. 5 years waiting period.	Does health status allow 3 hours of work per day in any job?	100% 2000: €731 2005: €730	Similar to early retirement pension. Assuming applicant would have earned last wage until 60. Actuarial deduction of 3.6% for each life year of receipt before 63 up to 10.3%	Available throughout entire time period for all cohorts.
Occupational DI (ODI)	Work disability in last or trained occupation	Social contributions paid in last 3/5 years. 5 years waiting period.	Does health status allow 6 hours of work per day in previous/trained occupation?	50% *(same as partial WDI post-2000[1])* 2000: €587 2005: €515	Same as WDI but is supposed to solely compensate for partial work capacity loss. About 12% of gross wage with average age at first receipt of 47, see Appendix B for details.	Cut for cohorts born after 1960. Effective insurance value higher, the higher wage in last occupation. WDI and ODI converge for low-income jobs

[1] Work capacity between 3 and 6 hours per day results in partial WDI at 50% of the benefits. Pre-2001, ODI benefit was 2/3 of WDI.

Source: own illustration. See main text for details. ODI was abolished for cohorts born after 1960 effective 2001. Appendix B provides details on the benefit calculation; changes in benefits affected all birth cohorts equally. Further, pre-2001, the health assessment applied an earnings threshold. The change to an "hours capacity assessment" also affected all cohorts equally.

Figure 2: Effect of 2001 Reform on Public DI Inflows Using Administrative Data

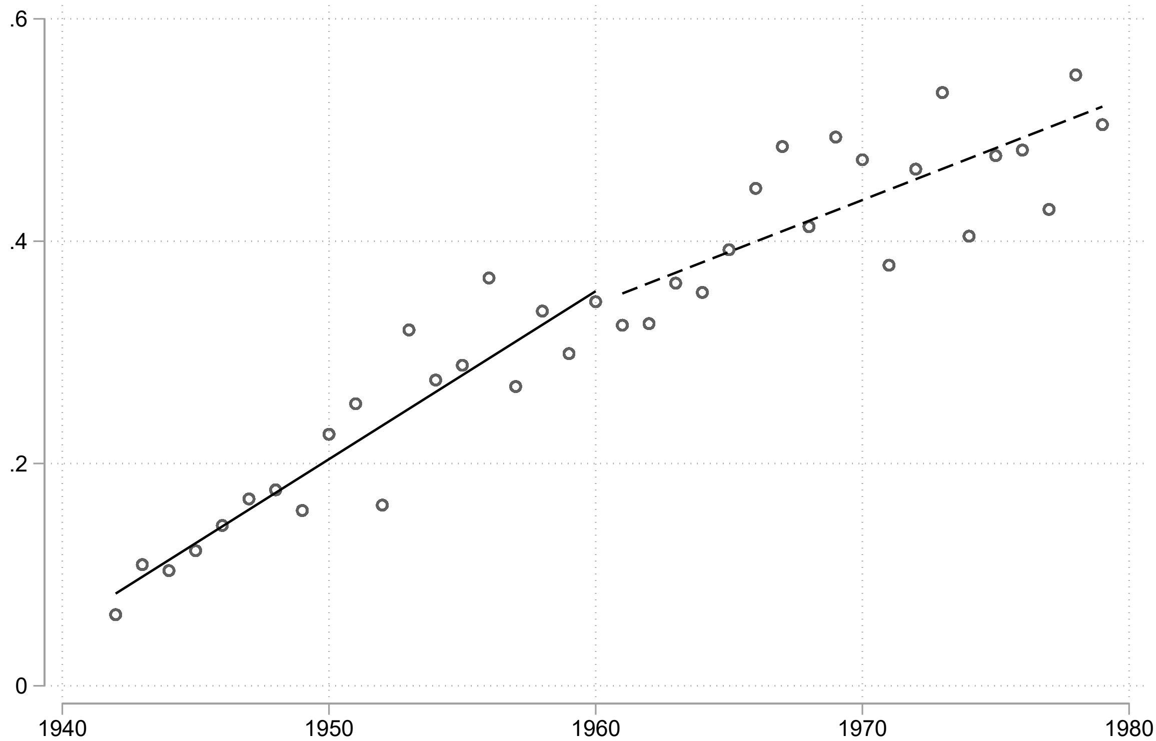


Source: Administrative SPI data on new public DI recipients by cohort and year. Treated cohorts are those born after 1960 and the treatment group; grandfathered cohorts are those born before 1961 and the control group. Figure plots $\beta D_c \times T_t$ estimates from equation 1 but with the post-reform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

Figure 3: Effect of 2001 Reform on Public DI Case Loads Using Representative SOEP Data

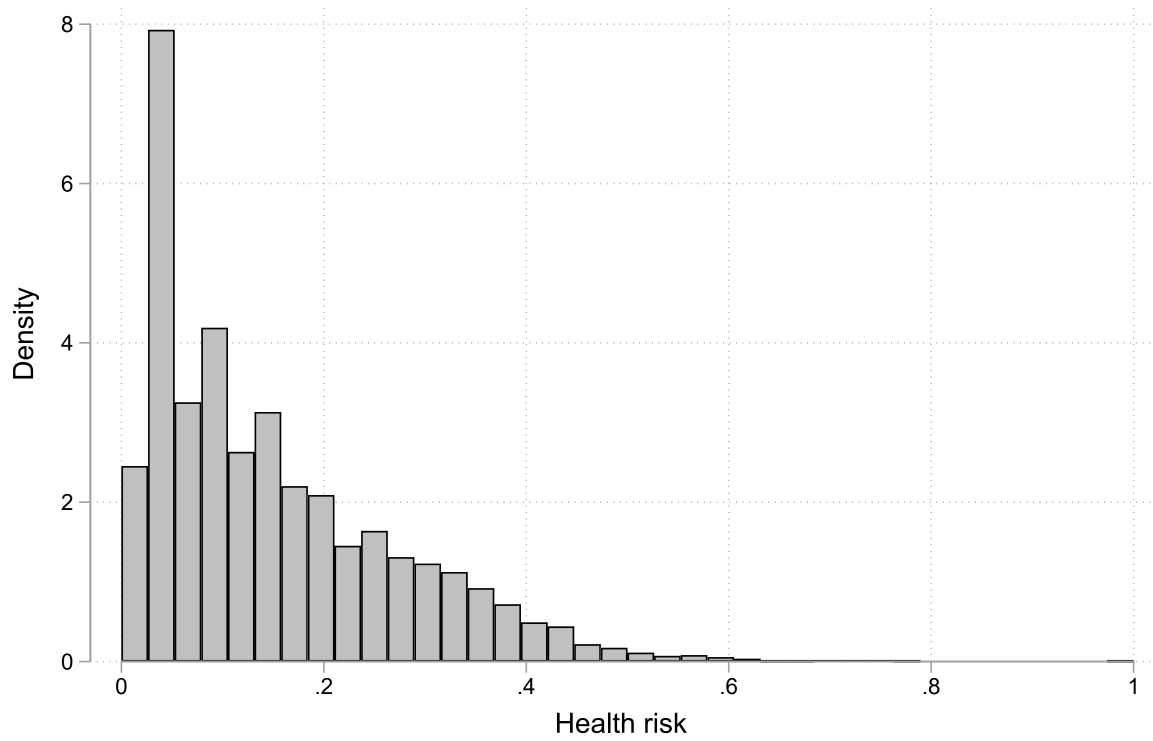
Source: SOEP v.33 – 95% sample. Sample is restricted to post-reform years. The figure is from one RD model similar to equation 2, estimated using quadratic trends in the running variable $f(z_i - c) = z_i - c$ to allow for different slopes before and after the cutoff. Robustness checks show results for 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 4: Effect of 2001 Reform on Private ODI Policies Using Representative SAVE Data



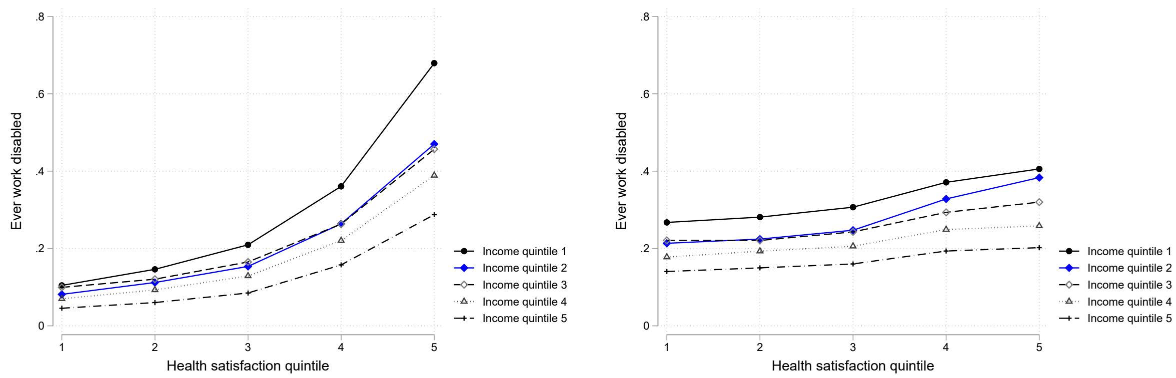
Source: SAVE data 2001-2010. The figure shows the raw nonparametric means of private ODI coverage by birth year, overlaid with separate linear trends before and after the cutoff. Other robustness checks vary the sample (Figure A8), vary the bandwidth (Figure A10, Calonico et al. 2020), study the smoothness 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 5: 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.

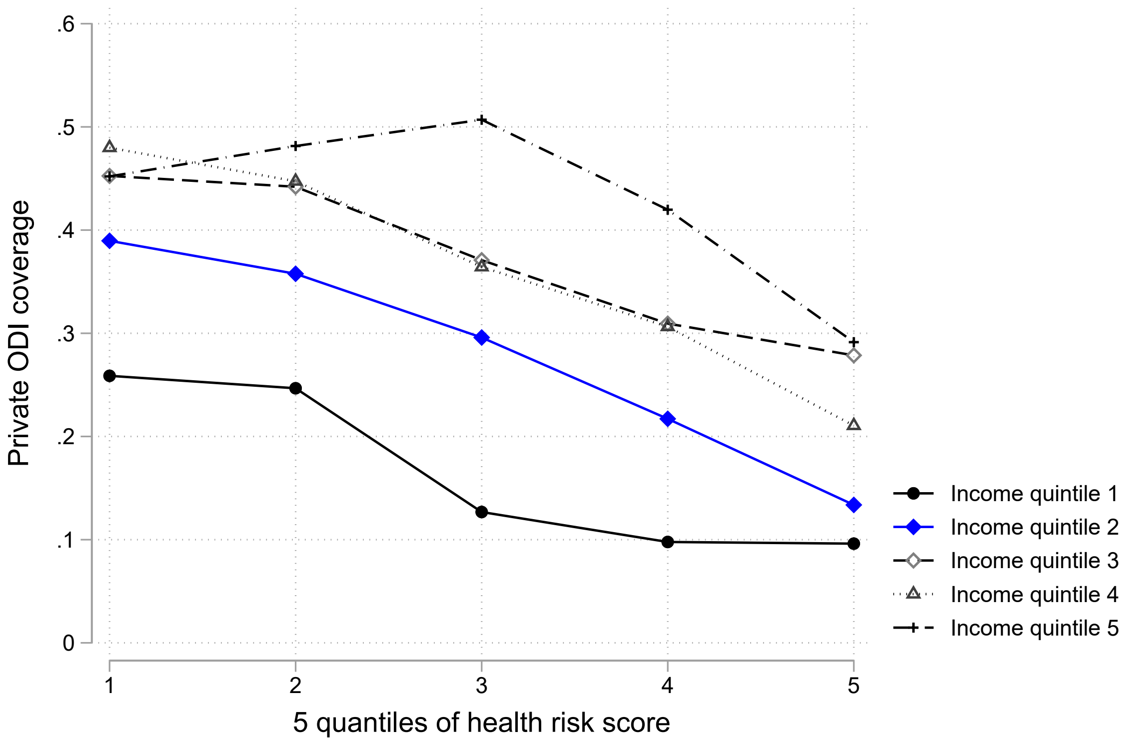
Figure 6: Lifecycle Risk of Work Disability by Income and Health Risk Score



(a) unconditional **(b)** conditional

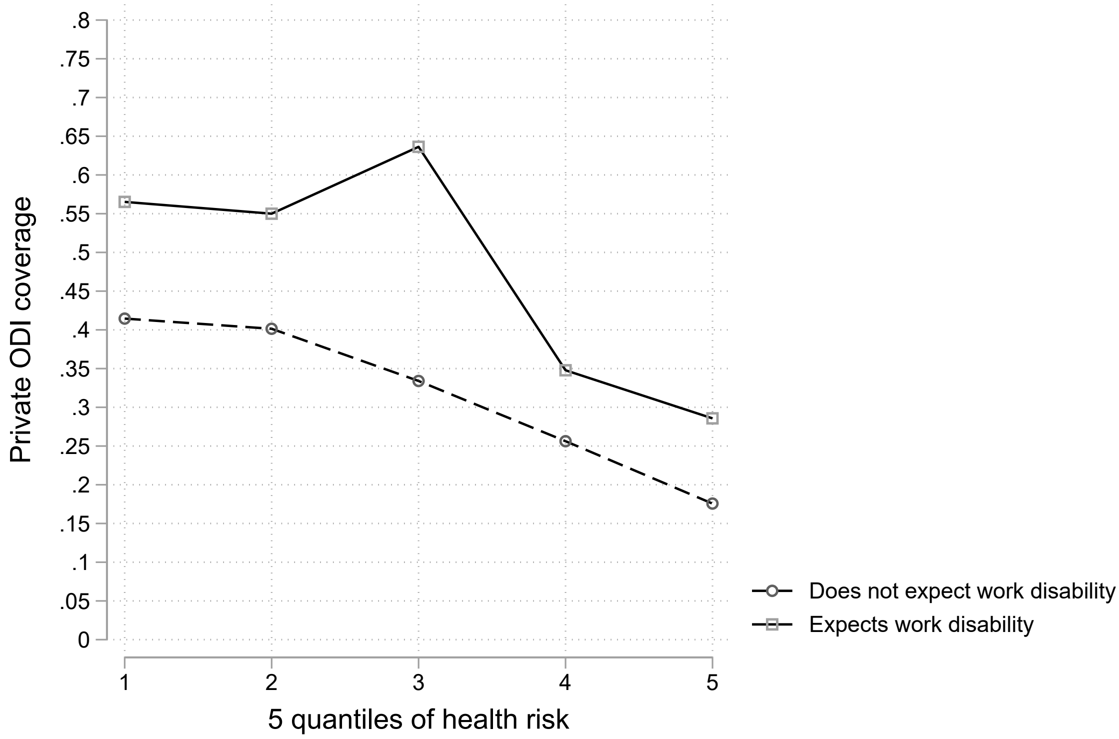
Source: SOEP v.33 – 95% sample. Figure 6a plots the unconditional risk of a severe health limitation over the working ages by the health satisfaction quintiles and the five net household income quintiles. Figure 6b first regresses the lifecycle risk of severe health limitations on socio-demographics, 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 7: Take-Up of Private ODI Policies by Health Risk and Income



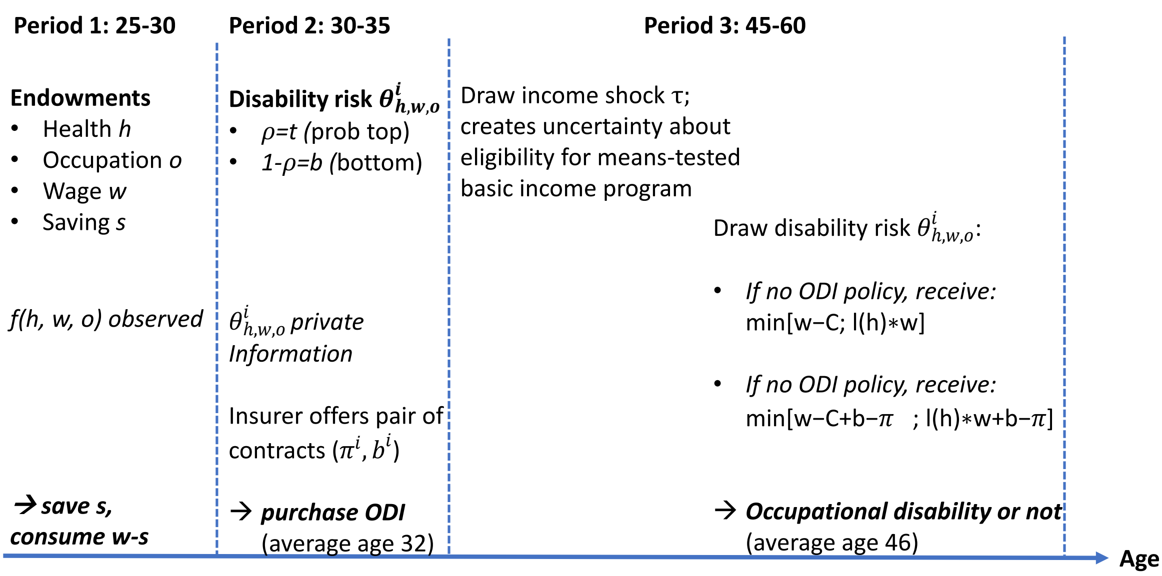
Source: SAVE data 2001-2010. Figure plots take-up rates of private ODI policies against the quintiles of the health risk score in Figure 5 and stratifies these curves by the five net household income quintiles.

Figure 8: Take-Up by Health Risk and Private Information



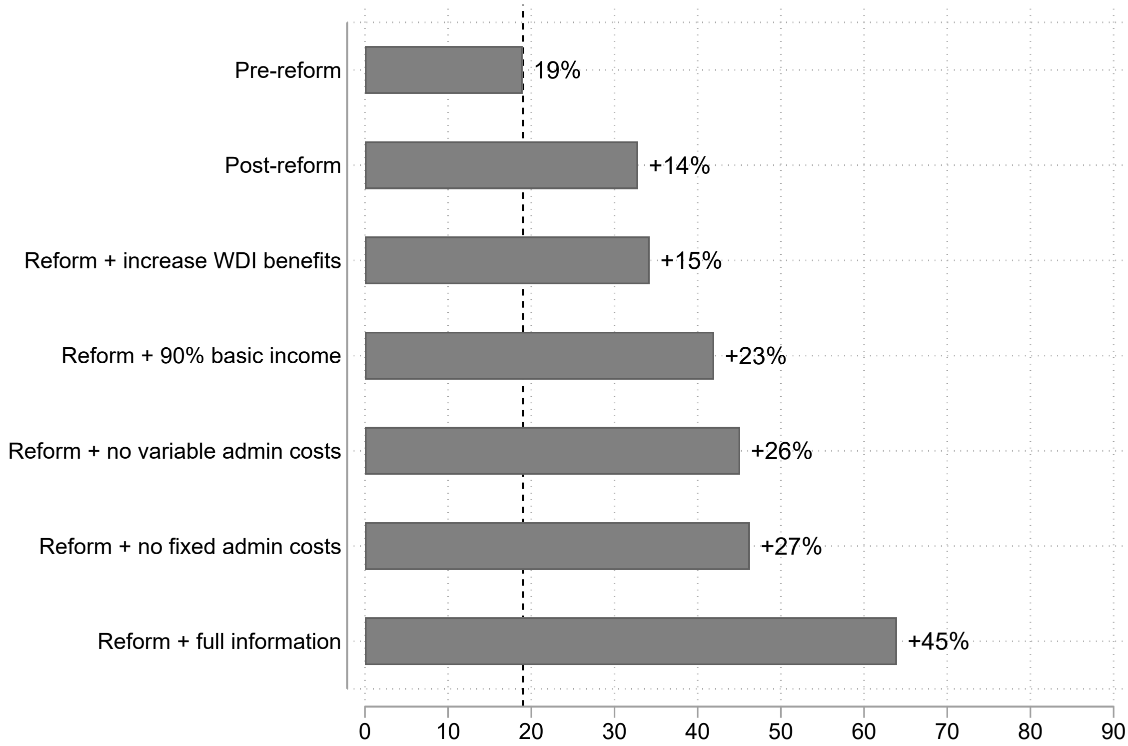
Source: SAVE data 2001-2010. Figure plots take-up rates of private ODI policies against the quintiles of the health risk score in Figure 5 and stratifies these curves by expected retirement before age 60. The latter information is directly elicited in the SAVE survey and proxies expected work disability.

Figure 9: Illustration of Lifecycle Time Periods in Baseline Model



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

Figure 10: Effect of 2001 Reform on Private ODI Policies by Health Risk Quintile



Source: The bars shows model predictions for average population-level reform effects. The private ODI take-up is simulated.

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
-0.085				
	(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	11,973	9,526	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	0.6222

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	0.2685	0.2736	0.1243	0.0975	0.1169
...	0.3963	0.3813	0.2996	0.1954	0.1535
...	0.4144	0.4515	0.3467	0.3070	0.2935
...	0.5037	0.4482	0.3778	0.3249	0.2685
Q5	0.4731	0.4564	0.5288	0.4411	0.3262

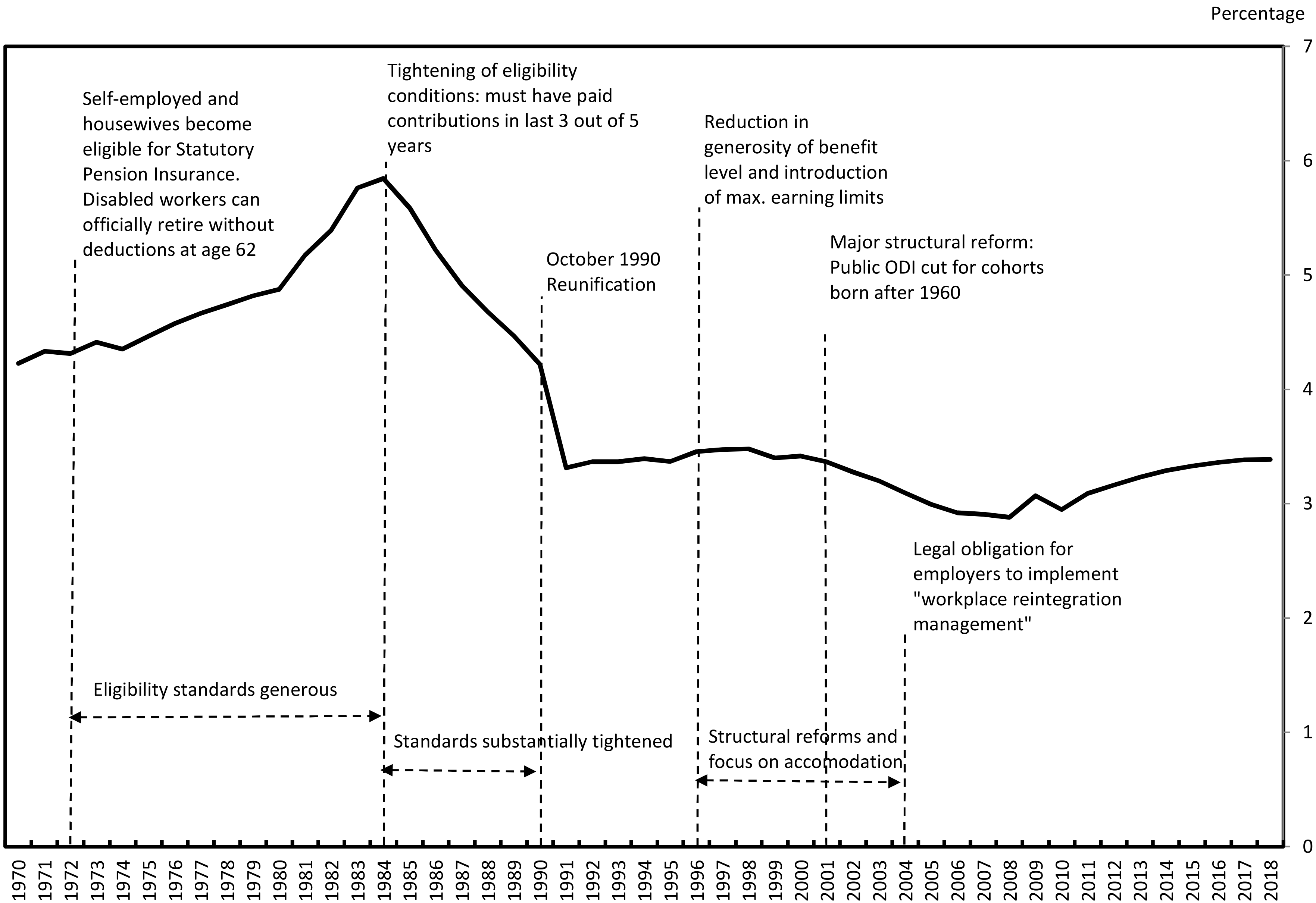
Table shows private ODI take-up rates by Health Risk (columns) and Income Quintiles (Rows). Q1 is the healthiest and poorest quintile, whereas Q5 is the sickest and richest quintile. Panel A shows the raw data from SAVE and Panel B show the private ODI take-up rates as produced by the general equilibrium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

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

	Pre-Reform	Post-Reform	+ 10% increase WDI	+ 90% basic income	No variable admin	No fixed admin	Full Info
Panel A: Total							
Take up	0.1894	0.3283	0.3422	0.4198	0.4510	0.4627	0.6394
Share of costs	0.8085	0.7313	0.7115	0.7754	0.8354	0.6838	0.9435
Load	0.2775	0.2890	0.2793	0.3126	0.1786	0.2729	0.5098
Profits	0.0118	0.0247	0.0230	0.0277	0.0337	0.0296	0.0535
Ex-ante utility	-1.4555	-1.4611	-1.4597	-1.4664	-1.4597	-1.4596	-1.4621
Govt total transfer	0.0726	0.0613	0.0606	0.0603	0.0603	0.0605	0.0613
DI transfer	0.0743	0.0576	0.0576	0.0576	0.0576	0.0576	0.0576
Basic income transfer	0.0038	0.0037	0.0030	0.0027	0.0027	0.0028	0.0037
Panel B: Good risks							
Take up	0.0283	0.1338	0.1508	0.2811	0.2505	0.2444	0.6845
Share of costs	0.2891	0.4026	0.4021	0.4109	0.4638	0.4452	0.9383
Load	0.2382	0.5855	0.6072	0.6607	0.6617	0.6412	0.6156
Panel C: Bad risks							
Take up	0.4548	0.6486	0.6574	0.6481	0.7812	0.8222	0.5651
Share of costs	0.8617	0.8430	0.8284	1.0359	1.0317	0.8006	0.9537
Load	0.6609	0.1883	0.1554	0.0638	-0.076 6	0.0926	0.2986

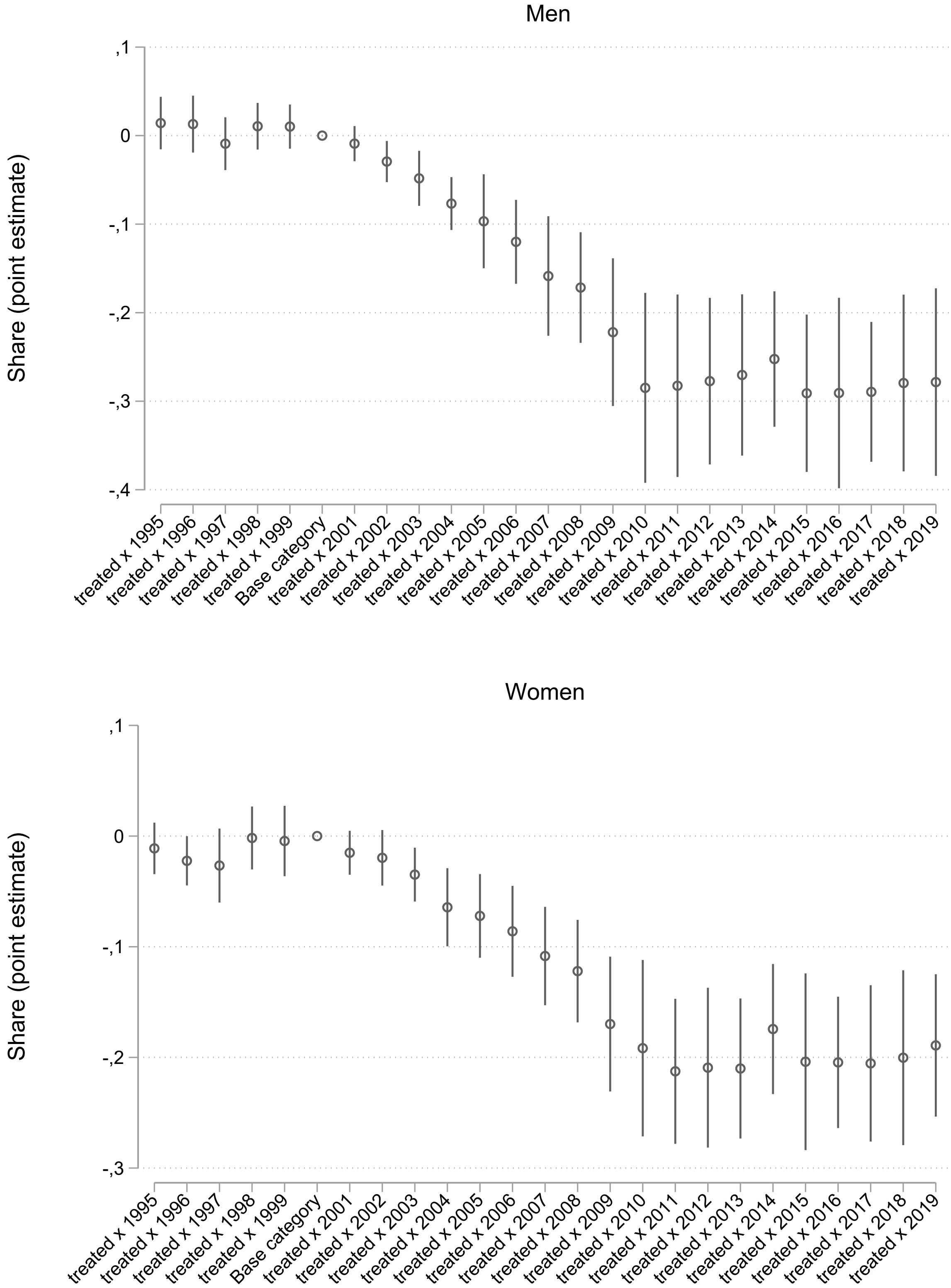
Table shows private ODI take-up rates, share of costs insured, loading factors, insurer profits, consumer utility and transfers by scenarios and good and bad risk types. All policies in columns (3) to (7) are combined with the 2001 ODI reform.

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



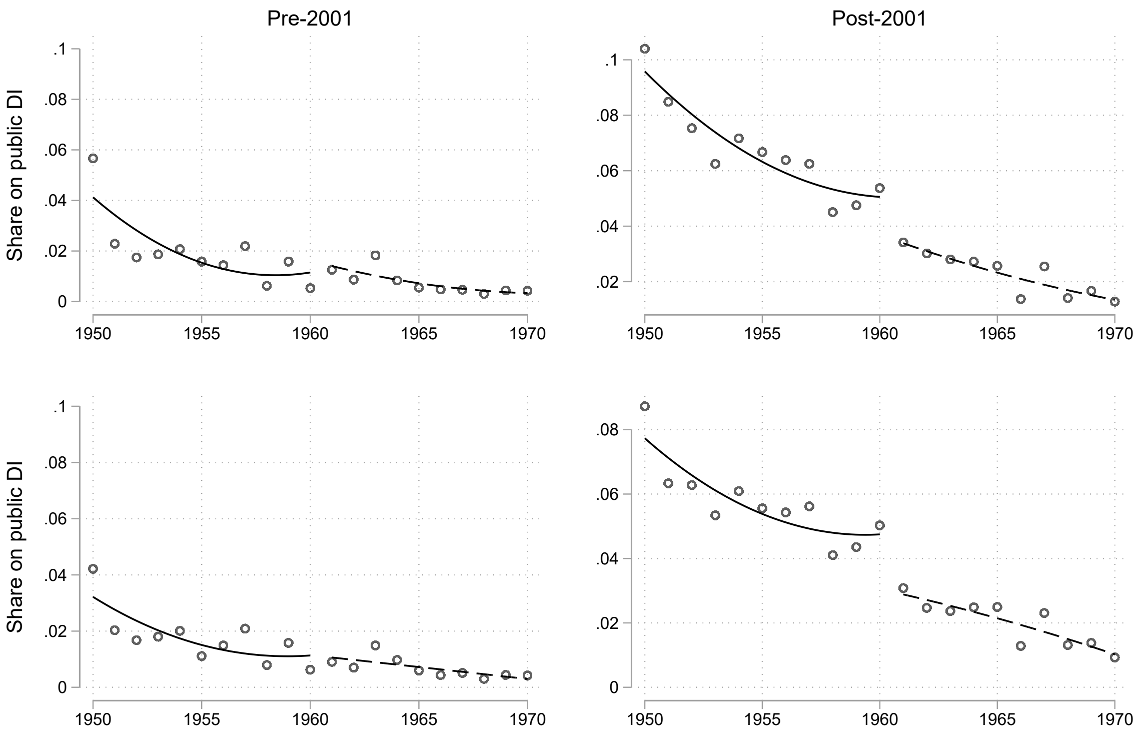
Source: Figure adapted from McVicar et al. (2022).

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



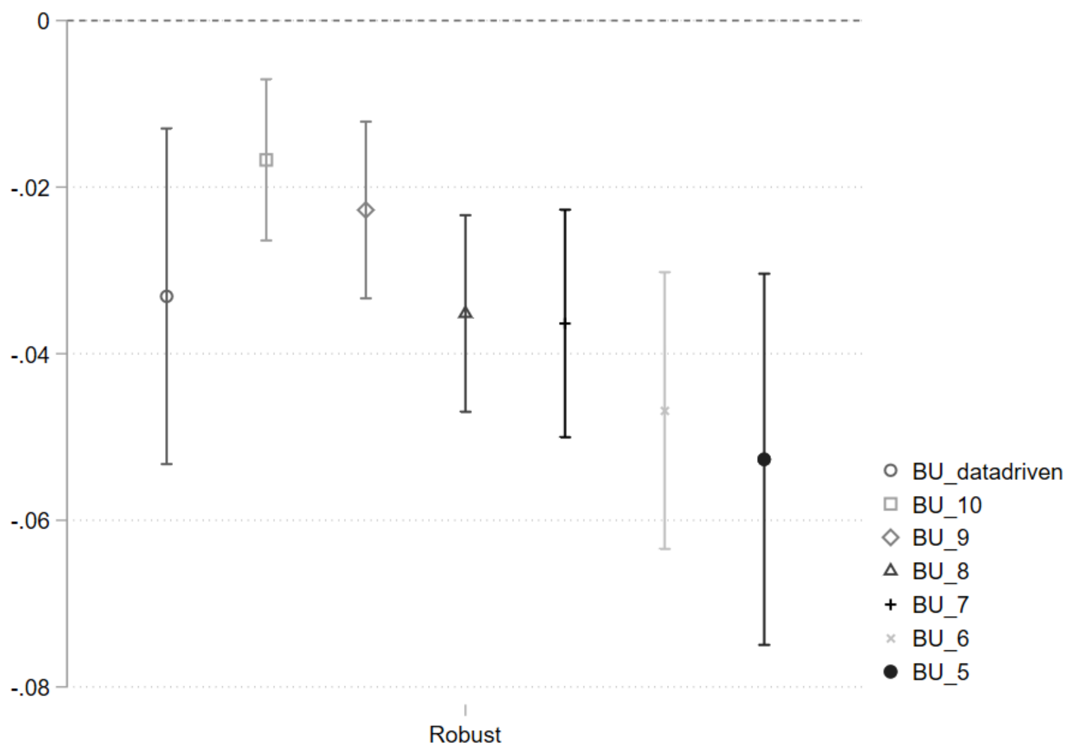
Source: Administrative SPI data on new public DI recipients by cohort and year. Treated cohorts are those born after 1960 and the treatment group; grandfathered cohorts are those born before 1961 and the control group. Figure plots $\beta D_c \times T_t$ estimates from equation 1 but with the post-reform 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

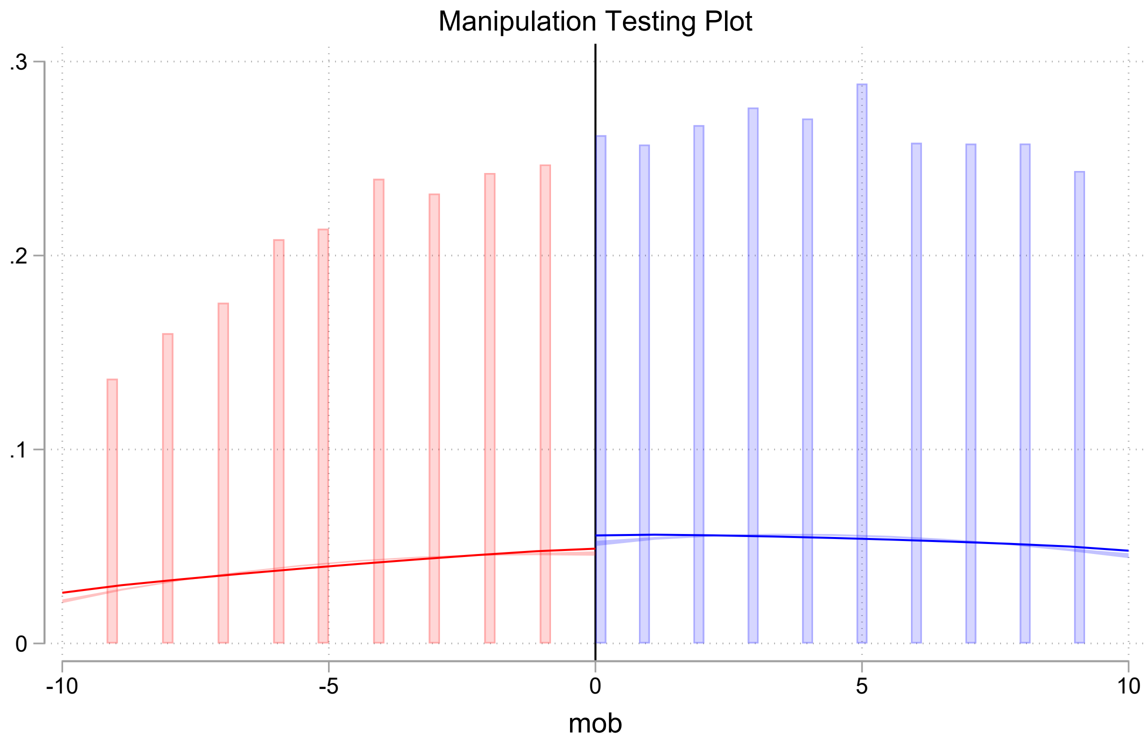


Source: SOEP v.33 – 95% sample. The figures show point estimates of robustness checks varying the bandwidths of RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks 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

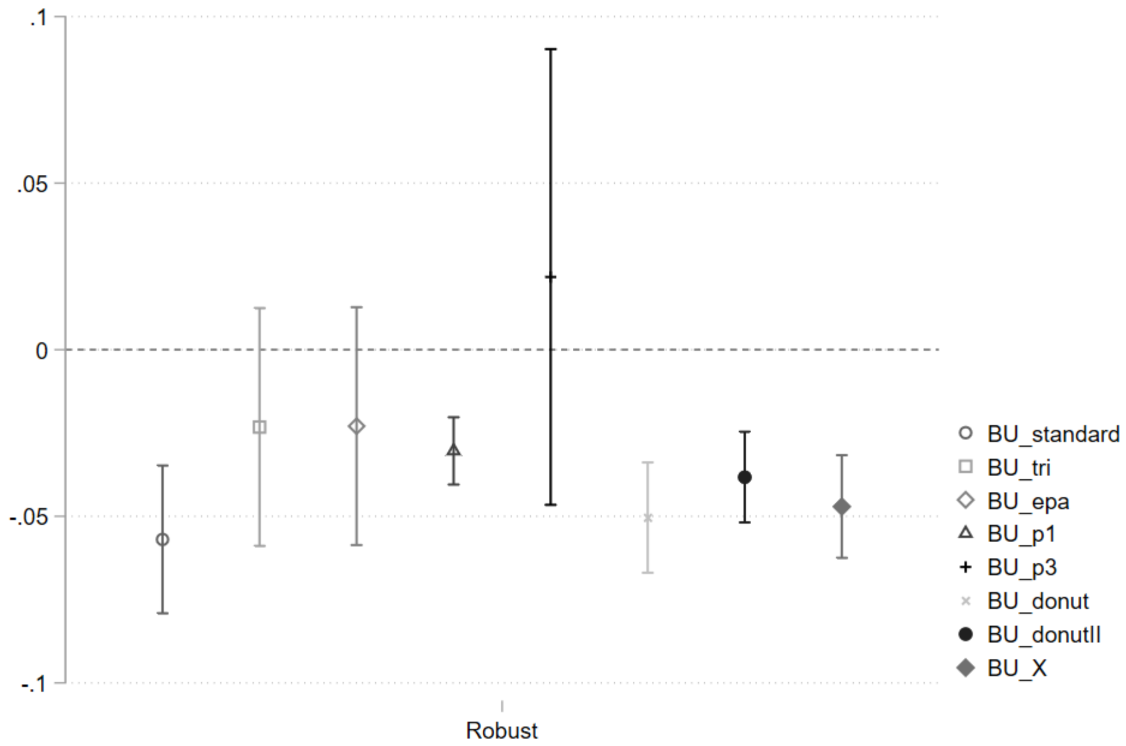
Source: SOEP v.33 – 95% sample. The figures show the raw nonparametric means of covariates as indicated, by birth year, overlaid with separate quadratic trends before and after the cutoff. Other robustness checks vary the sample and indicator (Figure A3), vary the bandwidth (Figure A4, Calonico et al. 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

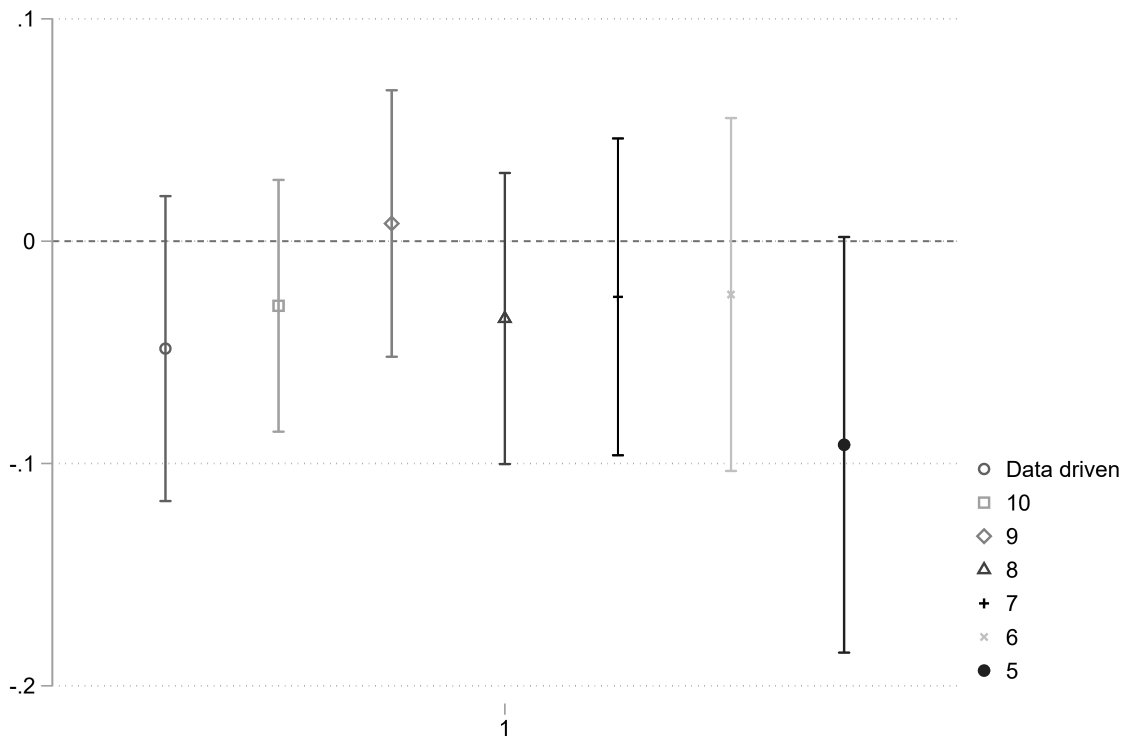


Source: SOEP v.33 – 95% sample. The figure shows the point estimates of a robustness check varying the order of the polynomials, varying weights, adding covariates,m and running donut RD models similar to equation (2), estimated using local polynomial regressions (Calonico et al. 2014, 2017, 2018, 2019). 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)

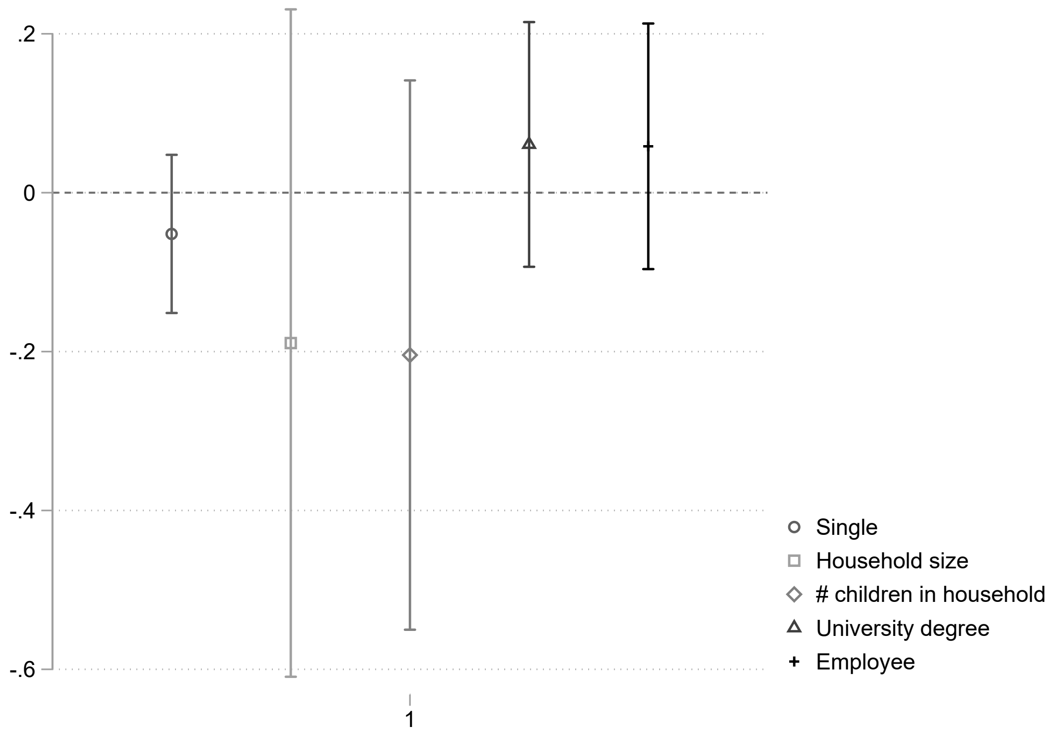
Source: SAVE data 2001-2010. The figures show the raw nonparametric means of private ODI coverage by birth year, overlaid with separate linear trends before and after the cutoff. The upper left graph is the default Figure (4), the upper right figure focuses on those eligible for Public DI, the bottom 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



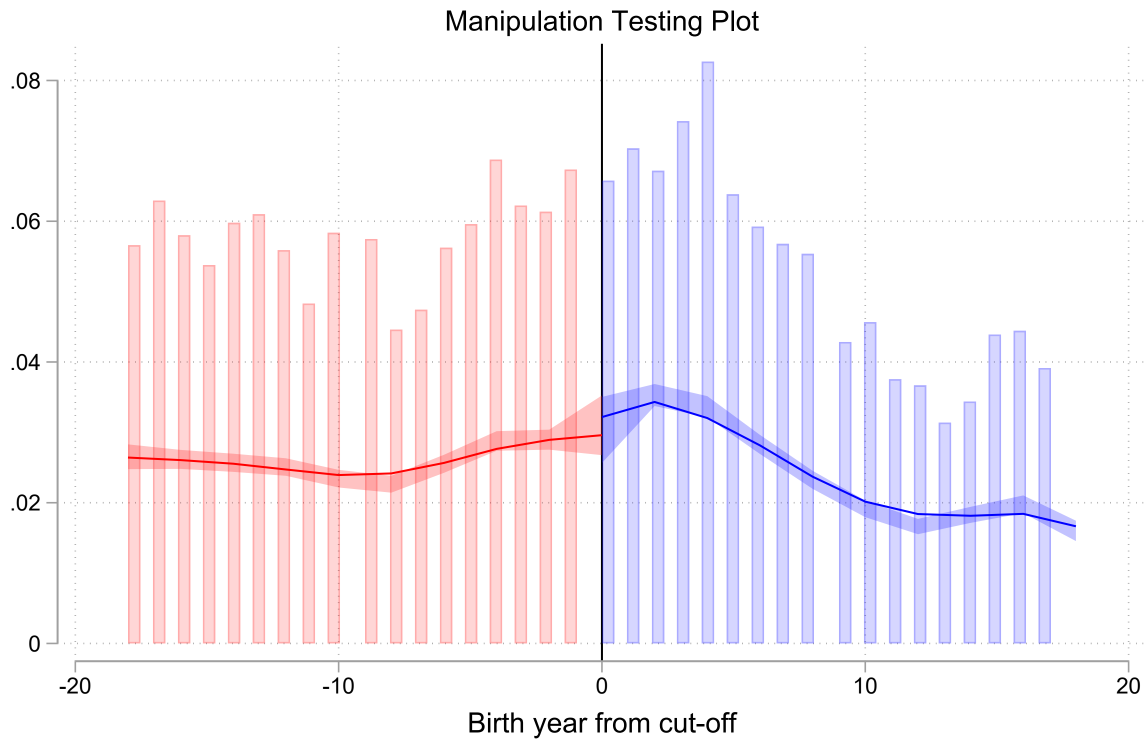
Source: SAVE data 2001-2010. The figures show point estimates of robustness checks varying the bandwidths of RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary 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



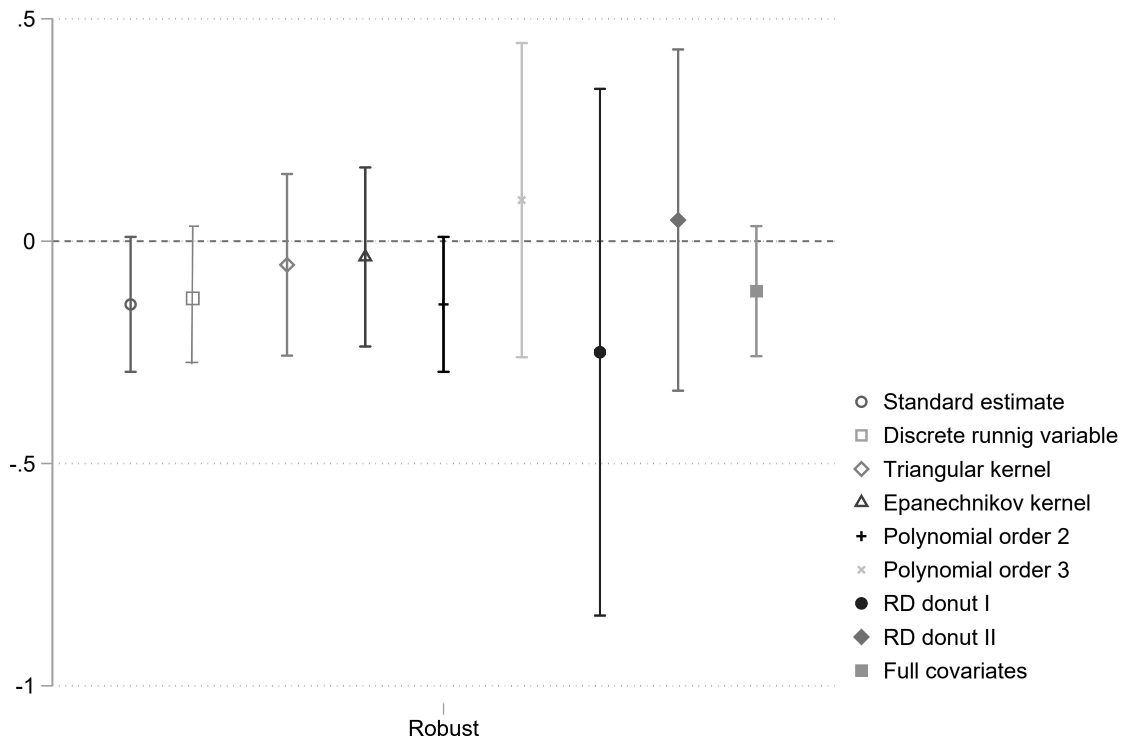
Source: SAVE data 2001-2010. The figures show point estimates of robustness checks testing for discontinuities in covariates using RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other 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



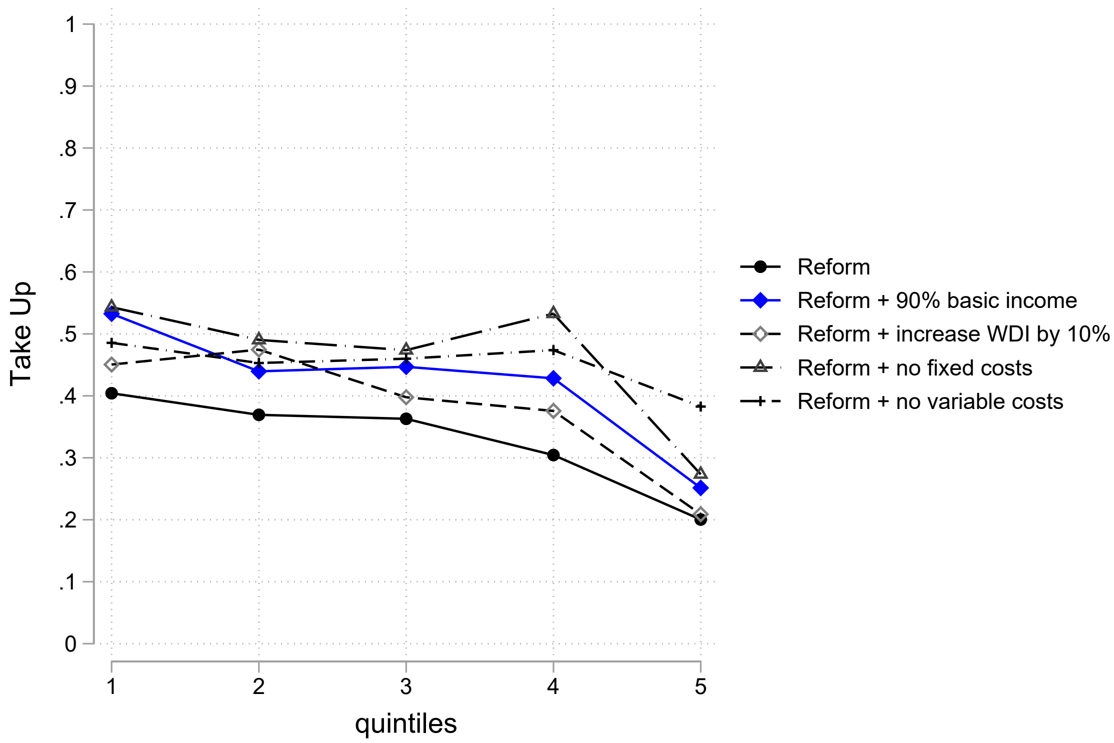
Source: : SAVE data 2001-2010. The figures shows a density plot of the running variable for RD models similar to equation (2), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al. 2014, 2017, 2018). Other robustness checks vary the sample (Figure 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: Take-Up Rates by Health Risk Score: 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 5. The other lines show take-up rates for alternative policy simulations by health risk quintiles using the general equilibrium model (see Section 5).

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: 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 A3), vary the bandwidth (Figure A4), study the smoothness of covariates (Figure A5), carry out density plots of running variables (Figure A6), and vary polynomials as well as carry out donut RDs (Figure A7).

Table A4: 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 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×	-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 5 by income quintiles. The first row shows the empirical moments form SAVE and the second row those produced by the model.

Appendix B: Benefit Calculation

We illustrate the effects of the 2001 pension reform on benefits by running a simple simulation 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 60.[29] These deductions are capped at 36 months or 10.8% ($AD_i = 0.892$).

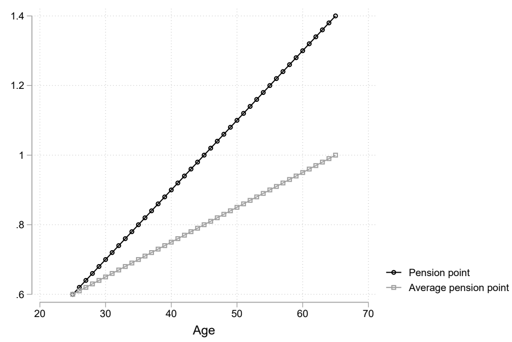
[29] In the meantime, the reference age has further increased to 63.

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

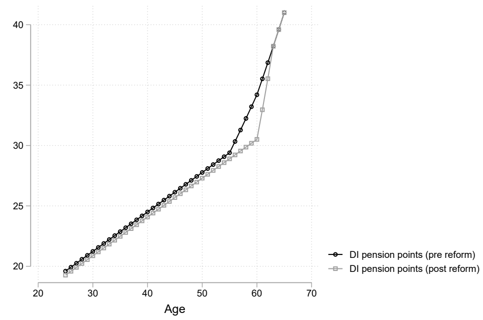
Simulation. Next, we simulate the effects of the 2001 reform on benefits for a stylized individual. We assume an increasing relative wage position that approximately equals 1 over the lifecycle. The individual starts working at age 25 and earns 60% of the average wage ($pp_{it} = 0.6$). The wage position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

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



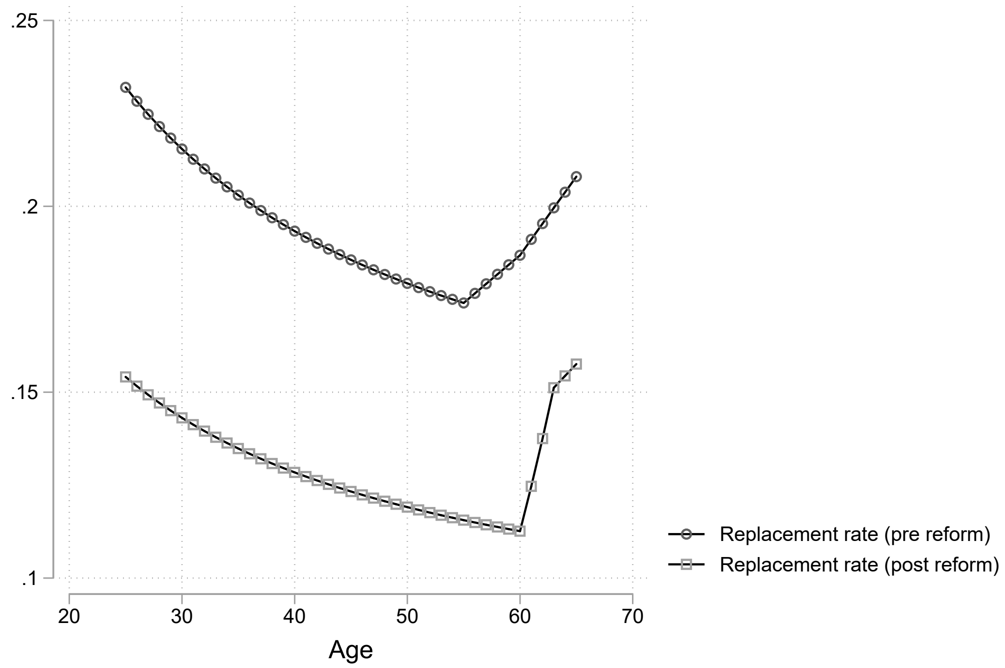
Source: own illustration. Note that the post-reform benefits apply either to the grandfathered cohorts who can still claim ODI benefits or the newly introduced partial DI benefits for people who are able to work more than 3 but less than 6 hours a day in any job.

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 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 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).

[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 treated cohorts are ineligible for ODI 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, we 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.

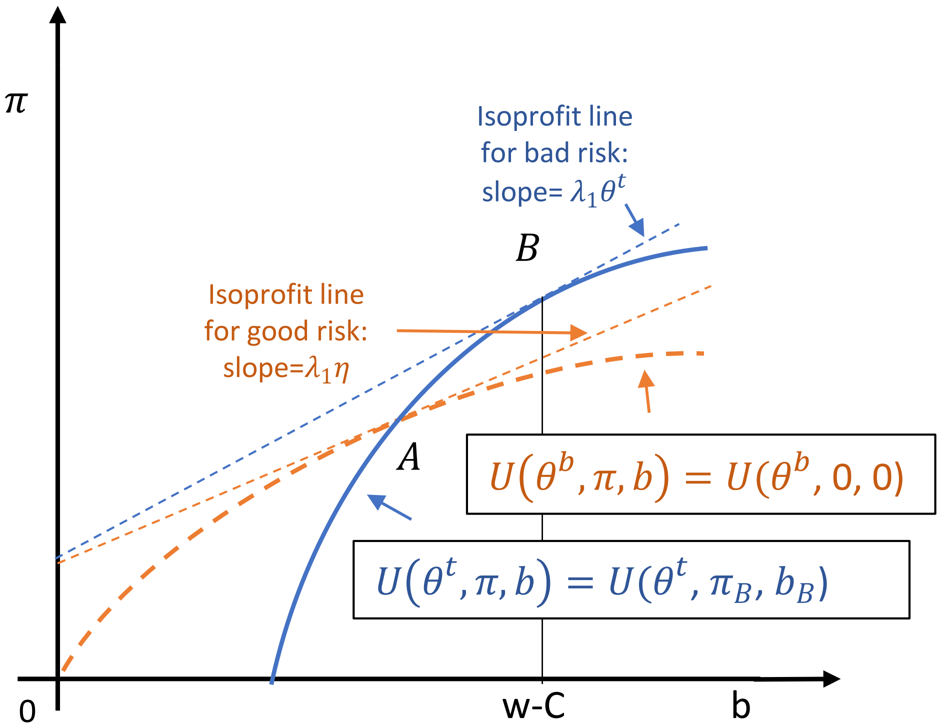
C1 Standard Case: Just Private Information

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.

Figure C1: Standard Case of Optimal Contracts with Private Information: Separating Equilibria



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.

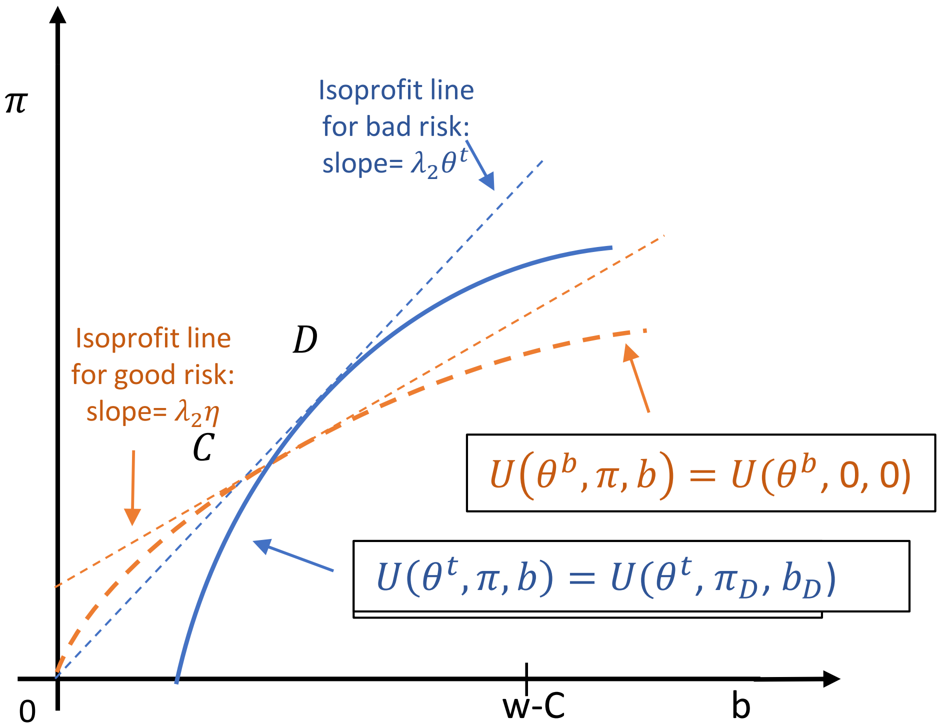
As discussed, the standard case cannot produce coverage denials by insurers. Only the good risks can be voluntarily uninsured with (0,0) and produce an ODI take-up that is not 100%. In other words, insurers always offer policies. Such as scenario can happen when the share of the population with low occupational 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

Figure C2: Optimal Contracts with Private Information and Admin Costs: Separating Equilibria



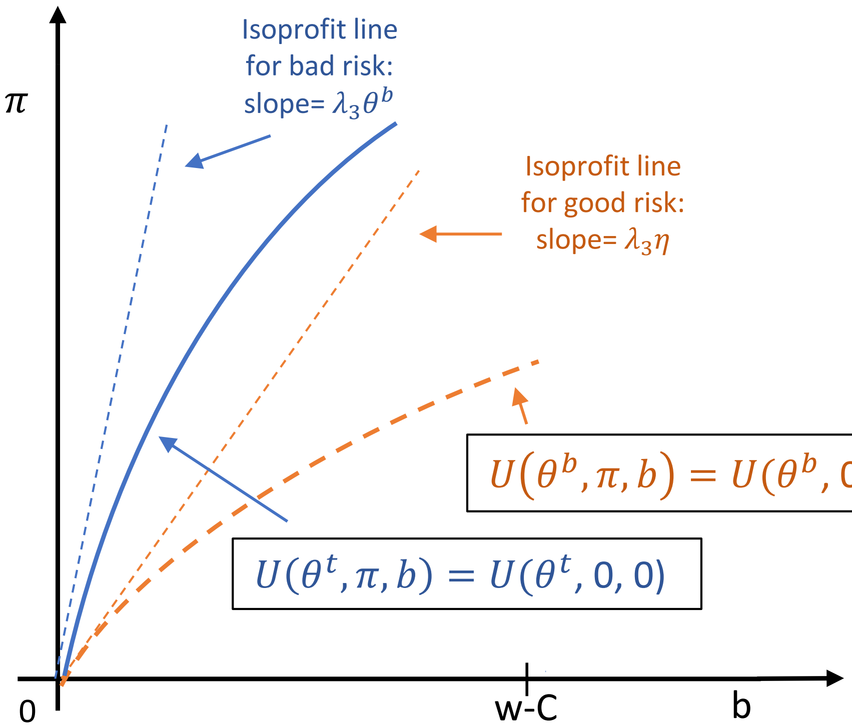
Source:: The dashed indifference curve shows optimal contracts for good risks at the bottom of the work disability distribution θ^b trading off premia (Π) on the y-axis and coverage levels (b) on the x-axis. The solid indifference curve shows optimal contracts for bad risks at the top of the work disability 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.

Figure C3: Optimal Contracts with Private Information and High Admin Costs: Denial



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 with 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 to 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 affects 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 9 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.