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# Fundamentally Reforming the DI System: Evidence from Germany

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**Abstract**

In 2001, Germany abolished public occupational disability insurance (ODI)—the second tier of its public DI system—for cohorts born after 1960. Using administrative data, we first document that, in the long run, overall DI inflows declined by roughly one-third. Second, using representative survey data, we document at best modest ODI insurance take-up responses in the private individual, risk-rated market, which lacks guaranteed issue. Third, an equilibrium model incorporating interactions between the public safety net, the first-tier public DI, and the private market reveals that coverage denials and weak insurance demand, driven by complementary social insurance, can explain the modest private ODI take-up response. Coverage gradients by income and health are thus substantial. Finally, counterfactual simulations highlight the limited scope of incremental reforms.

**Keywords:** occupational disability insurance, individual private DI, coverage denials, risk rating, private information, adverse selection, social safety net

**JEL Classifications:** D14, D82, H53, H55, I14, I18, I38, J14, J26

Please see Appendix Section A for a complete list of acknowledgments. Please note that "Technical Reports and Web Links" are at the end of the Online Appendix.

## 1 Introduction

For decades, the design of social insurance systems has been a central focus of economic research (cf. Aizawa et al., 2025). Three core pillars exist across all OECD (2010) countries: unemployment insurance, workers' compensation, and disability insurance (DI). Hence, evidence from one OECD country may offer valuable lessons for others (Besharov and Call, 2022).

A rich literature in economics has examined the implications of public disability insurance for labor supply (Deshpande, 2016), earnings and consumption (Ruh and Staubli, 2019), beneficiaries' health and well-being (Gelber et al., 2023), intergenerational welfare dependence (Dahl et al., 2014), and firm accommodation (Aizawa et al., 2025). While Dutch reforms introducing experience-rated premiums for employers have received considerable attention (cf. Koning and Lindeboom, 2015), much less is known about the 2001 German DI reform, which reduced program generosity and marked a shift toward a mixed public–private insurance system. This is noteworthy, as U.S. reform proposals explicitly reference the German and Dutch reforms, arguing that private markets could offset reduced public DI generosity (cf. Autor and Duggan, 2010; Burkhauser and Daly, 2022).

This paper examines the response of the private disability insurance (DI) market to the 2001 German reform, which eliminated public occupational disability insurance (ODI) for cohorts born after 1960. ODI insures the risk of work disability in one's *occupation* or a comparable job. As Germany has maintained a co-existing public work disability insurance (WDI) program covering the risk of work disability in *any* job in the economy—alongside a means-tested cash transfer program—public ODI primarily insured higher-income employees against income losses up to the consumption floor. Consequently, public ODI's purpose was commonly referred to as "social status protection." The preamble to the reform bill, introduced by the Social Democratic–Green coalition, emphasized this motivation:[1]

> "There is widespread agreement among experts and academics, as well as in the political arena, that a DI reform is needed. [...] In particular, public ODI has been criticized as it has become a 'prestige pension' for insured individuals with particular qualifications in high-level positions." (Bundestag, 2000)

**[Insert Figure 1 about here]**

Figure 1 illustrates the eligibility principles of ODI and WDI, as developed by the *German Federal Social Court*. Six occupational categories range from managerial positions at the top to unskilled work at the bottom of the labor market. ODI eligibility is established when a disability prevents a worker from performing their previous job, or one within the same or the next-lower category. WDI eligibility, by contrast, requires the inability to perform *any* job across all categories. The insurance value of ODI, therefore, differs structurally across these categories. For low-income earners in basic jobs, being unable to work their previous job generally implies being unable to work in any job. For higher-income earners, by contrast, public ODI functioned as a publicly funded supplemental insurance, preserving their social status.

[1] A prior draft by the three center-right parties made a similar point: "As a consequence, public ODI pensions are a privilege of insurees with extraordinary education and in elevated positions. However, the insurance principle [...] requires that insurees [...] must have equal opportunities to claim benefits" (Bundestag, 1997).

As these institutional features are essential to understanding private ODI take-up following the abolition of public ODI, the theoretical framework developed in this paper models the interactions between public WDI, public ODI, and the private ODI market from both the insurer's and the consumer's perspectives.

Before introducing the theoretical framework, however, we first show that the reform had a tangible impact on the treated cohorts. Using administrative data spanning 18 post-reform years, we document how the reform affected overall DI enrollment. Event-study estimates show that abolishing public ODI led to a reduction of more than 30% in total DI inflows (WDI + ODI) in the long run, approximately ten years after the reform's implementation.

The paper's main part then analyzes how the private ODI market responded using reduced-form and structural equilibrium approaches. Based on representative survey data and a Regression Discontinuity (RD) design with birth year as running variable, we find at best modest increases in private ODI take-up, consistent with concurrent research using data from a single private insurer (Seibold et al., 2025). Our 95% confidence intervals provide no evidence that coverage in the general population exceeded 55% after the reform.

Why has private ODI take-up not increased more strongly? To better understand the underlying economic mechanisms, we employ a general equilibrium model that captures key demand and supply features of the German private DI market. We build on Braun et al. (2019), tailoring their framework to German market regulation and social insurance interactions. The model features an optimal contracting framework that endogenizes insurers' supply decisions and the design of private ODI policies. Specifically, insurers offer two contracts for each of 750 risk groups, one for high-risk and one for low-risk types, where true types are unobserved. Pricing is based on observables and incorporates health risk rating, as in reality.

Another crucial market feature captured by the model is coverage denials. As in the U.S. pre-ACA health insurance market, German applicants must disclose all current and past medical conditions and their use of healthcare. This information is stored in a shared database accessible to all insurers, generating strong selection into applications (Finanztip, 2017; informa HIS GmbH, 2025). Rating agency data covering nearly the full market indicate that, conditional on applying, about one-sixth of applicants are entirely or partially denied coverage, for example, via preexisting-condition clauses. Yet these figures reflect only *official* applications. A leading consumer magazine cites market insiders with roughly 235,000 annual coverage denials, implying rejection rates near 50%.[2] Moreover, an audit study of 110 anonymous applications across 59 policies found that 81% of applicants were offered less generous coverage than requested, and often denied coverage entirely, even for minor health issues (Ökotest, 2014). One of the model's key features, therefore, aligns closely with this empirical reality: insurers routinely deny coverage based on observable characteristics.

To our knowledge, this is the first paper to model disability insurance supply and demand within an equilibrium framework that endogenizes private policy offerings and allows for both risk rating and coverage denials. The model builds on the classical optimal contracting framework of Rothschild and Stiglitz (1976) and, specifically, heavily on Braun et al. (2019), adapted to the regulation and structure of the German private ODI market. It provides a framework for understanding the modest private ODI take-up responses and replicates the strong income and health gradients observed in the data. The muted take-up responses can be explained by the relatively generous means-tested cash transfer program and public WDI, together with coverage denials and risk rating. Recall that ODI offers "social status protection" against income losses due to health shocks (Figure 1). Accordingly, we observe crowding out of private ODI coverage by public ODI primarily among higher-income and healthier individuals–those for whom ODI yields a high(er) insurance value, lower premiums, and fewer coverage denials. The final section simulates potential complementary reforms, showing that private ODI coverage remains very incomplete under all policy scenarios.

In complementary work, Seibold et al. (2025) also study private ODI take-up following the 2001 reform. Using data from a single private insurer, and consistent with our representative evidence, they find modest increases in private ODI sales and a low observed willingness to pay. However, they employ different data, assumptions, and a distinct modeling approach—an adaptation of the Einav et al. (2010) sufficient-statistics framework (Kleven, 2020). They argue that distributional concerns could justify public ODI provision. While this paper and Seibold et al. (2025) share several consistent findings, they also differ in many important dimensions. This paper contributes to the literature beyond Seibold et al. (2025) as follows:

First, it uses representative, rich survey data rather than data from a single insurer. In particular, the analysis draws on two representative panel datasets, SAVE (Saving for Old Age in Germany) and SOEP (German Socio-Economic Panel Study). We use that data to estimate

[2] This number is plausible because applicants routinely request anonymous "pre-assessments" before formally applying and having their information stored in the HIS database (Verbraucherzentrale, 2024).

reduced-form models and extract key empirical parameters to inform our model, such as private ODI take-up by income and health status, and the probability of work incapacity by income and health status.

Second, these empirical parameters are fed into a general equilibrium model that explicitly captures the supply and demand sides of the private ODI market, as discussed above. Our model adopts a long-run perspective and generates equilibrium outcomes. In contrast, Seibold et al. (2025) use a static, reduced-form framework based on sufficient statistics designed to evaluate marginal policy changes. Their approach treats the supply side as given and abstracts from coverage denials and co-existing social insurance strands.

Third, on the supply side, the model endogenizes contract design, including health risk rating and coverage denials—both critical for understanding how the market responded to the reform. Coverage denials, in particular, are crucial in explaining the modest take-up responses and the socio-demographic gradients in take-up.

Fourth, on the demand side, the model approximates how individuals decide whether to purchase private ODI when entering the labor market, accounting for uncertainty in lifetime income and lifecycle work incapacity risk, the public WDI system, and the means-tested safety net. We estimate these key demand-side parameters for our structural model using the long-running and representative SOEP and SAVE datasets.

Fifth, we explicitly model the interaction between public WDI and private ODI, and how these programs, together with Germany's means-tested cash transfer system, jointly shape private ODI demand.

Finally, unlike Seibold et al. (2025), we find empirical evidence of adverse selection and therefore model private information on the demand side. Hendren (2013) shows, both theoretically and empirically, that high-risk individuals often hold private information, leading to coverage denials—an important mechanism consistent with our framework.

Beyond the specific contribution relative to Seibold et al. (2025), this paper contributes to a small but important body of international economics literature on interaction effects between social and private insurance (Borghans et al., 2014; Autor et al., 2019; Leung and O'Leary, 2020; Ahammer et al., 2025). While the literature on public and private health insurance has demonstrated a crowd-out effect of private through public insurance (Cutler and Gruber, 1996; Clemens, 2015), the DI literature is much thinner, potentially due to institutional differences across countries. Although individual private DI exists in Canada and the United States (Autor et al., 2014), their markets for private *group* DI policies, provided by employers, is much bigger.[3] Autor et al. (2014) study determinants of private DI spells in the U.S. group market, whereas Cabral and Cullen (2019) find that U.S. employees value public DI more than twice its costs. Public DI has also been found to insure more than just health risks (Deshpande and Li, 2019), similar to health insurance (Lockwood, 2025).

This paper also contributes to a rich literature on selection and supply-side competition in insurance markets (Fang et al., 2008; Aizawa and Kim, 2018). Numerous papers have analyzed insurer-provider interactions in U.S. health insurance (Shepard, 2022) and the relationship between public and private reimbursement schemes (Clemens and Gottlieb, 2017; Carey, 2017; Lavetti and Simon, 2018). While these studies highlight the importance of the supply side in insurance markets, private DI markets more closely resemble life insurance (Fang and Wu, 2020) and German long-term health insurance (Atal et al., 2025), which are purely financial contracts where insurer-provider negotiations generally play little role.

Naturally, this paper also contributes to the literature on German DI. Existing research has described the German private ODI market (Soika, 2018; McVicar et al., 2022) and examined aspects of the 2001 reform. For example, Börsch-Supan et al. (2022) find no evidence that the reform improved target efficiency, and Hanel (2012) find little effect of lower WDI benefit levels on DI inflows. Although not evaluating the 2001 reform, Seitz (2024) uses the same claims data as Seibold et al. (2025) to estimate a dynamic lifecycle model, concluding that in the presence of a private market, the welfare-maximizing public DI program would be less generous than in its absence, a finding consistent with our framework.

Finally, the paper contributes to the international literature on DI. Dahl et al. (2014) study a Norwegian reform that allowed DI recipients to retain a larger share of earnings, showing that it increased labor supply and reduced program costs, consistent 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 eligibility rules outperform benefit cuts. Other studies demonstrate that barriers to applying for DI, such as health screening, shape program inflows (Autor and Duggan, 2003; De Jong et al., 2011).

[3] This group market also includes *short-term* DI (also called "medical leave" or "Temporary Disability insurance, TDI"), 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.

# 2 The German Disability Insurance System

## 2.1 Social Insurance in Germany

Germany has a European-style, comprehensive social safety net for employees, consisting of five social insurance branches: statutory unemployment, accident, health, long-term care, and pension insurance (SPI). Eligibility for sick leave and medical benefits is universal and comparatively generous by international standards (Ziebarth and Karlsson, 2010, 2014).

Additionally, Germany operates a universal, means-tested cash transfer program that provides a de facto consumption floor. In 2025, monthly cash benefits for a childless adult amounted to €563 (about $600), *in addition* to rent and utility coverage.[4] Below, we examine the role of this program in explaining the low private ODI take-up rates among low-income individuals.

## 2.2 The Public Disability Insurance Reform of 2001

Germany's public DI program is part of SPI. Employers and employees each contribute a 9.3% payroll tax on monthly earnings up to a ceiling of €8,050 (in 2025). To qualify, employees must have paid SPI payroll taxes in at least three of the five years preceding the application.

Before 2001, the public DI system comprised (a) work-disability insurance (WDI) and (b) occupational-disability insurance (ODI). The 2001 reform abolished public ODI for cohorts born after 1960 (the treated), while those born before 1961 were grandfathered (the controls) and retained eligibility. Appendix Figure B1 illustrates the core principles of both programs.

WDI insures against *general* work disability—the inability to perform *any* job in the economy. ODI, by contrast, insures against *occupational* disability, often referred to as "social status protection." Occupationally disabled individuals are those who *"[. . . ] due to health reasons, are unable to work in either their trained or a comparable occupation in terms of education and skills required"* (Deutsche Rentenversicherung, 2023a). As discussed, ODI's insurance value is structurally higher for high-income workers in specialized professions. For basic occupations, WDI and ODI effectively converge, since workers in low-skill jobs cannot usually transition to lower-paid work (Benen, 2023). Thus, by design, ODI primarily benefits higher-income employees.

[4] A 2004 reform introduced *Arbeitslosengeld II*, which decoupled means-tested benefits from prior labor income and shortened benefit duration. This reform did not differentially affect treatment and control cohorts of the 2001 DI reform.

**Benefit Calculation.** Public WDI and ODI benefits are calculated as ”early-retirement pensions” with actuarial reductions, assuming recipients would have earned their pre-disability income until age 60. Benefits depend on lifetime earnings histories and are not adjusted for household composition, income, or assets. Appendix B provides detailed simulations. Before the reform, simulated public ODI replacement rates at age 46—the average age of occupational disability—were approximately 18% (€ 587 in 2000)—compared with 32% for WDI. Public ODI benefits were never intended to provide full income replacement but rather to compensate for *partial* losses in work capacity when employees moved to lower-paid or part-time work.

**2001 Reform.** The reform eliminated public ODI for post-1960 cohorts and introduced several accompanying adjustments (Appendix B). These additional changes did not differentially affect birth cohorts: WDI benefits were reduced for all, and ODI benefits for grandfathered cohorts. Although Hanel (2012) finds little evidence of behavioral responses, the lower benefit levels may have reduced the attractiveness of applying for ODI. If so, our estimates of the reform's impact on total DI inflows represent a lower bound.

After outlining the structure of the public disability insurance system, we now turn to the private market, which operates independently and plays a complementary role for individuals seeking to insure against occupational disability risks.

### 2.3 Private Disability Insurance in Germany

**Basic Principles.** Germany's private occupational disability insurance (ODI) market is predominantly an *individual* market, in contrast to the group-based systems common in the United States (Autor et al., 2014). Similar to Germany's long-term private health insurance market (Atal et al., 2019, 2025), private ODI policies are individually underwritten. Initial premiums depend on age, medical history, health behaviors, occupation, and income, as reported in detailed health questionnaires (Appendix Figure B2). As a result, premiums for high-risk occupations can reach several hundred dollars per month, and coverage denials are common (Ökotest, 2014). However, premiums are only risk-rated once at inception and not dynamically adjusted after changes in health. Private ODI contracts fall under the *Insurance Contract Act* and represent bilateral agreements between insurer and policyholder with one-sided commitment. Similar "own-occupation DI" products exist in other countries, notably the United States, where they are marketed to high-income professionals such as physicians and lawyers (Brian SO Insurance, 2023).

**Age at Inception and Claiming.** The average age at policy inception is 32, the average age at disability onset is 46, and contracts typically extend to age 64 (Morgen & Morgen, 2021). These averages inform our theoretical framework for modeling the decision to purchase private ODI.

**Coverage Denials.** Rating-agency data covering nearly five million policies show that, in 2019, 23% of applicants were either denied coverage (8%), offered contracts with pre-existing condition clauses (11%), or charged additional risk premia (4%) due to medical histories (Morgen & Morgen, 2021). These shares are *conditional* on "officially" applying for a policy. In practice, brokers and online tools allow potential applicants to pre-assess eligibility. To avoid permanent storage of medical data in the industry's shared IHS database, applicants often request "anonymous pre-assessments" (Verbraucherzentrale, 2024). Industry estimates suggest roughly 235,000 annual rejections—an effective rejection rate near 50% (Morgen & Morgen, 2021). Moreover, an audit study using 120 fictitious applications found that 81% of applicants were offered less generous coverage than requested (Ökotest, 2014).

**Premiums and Benefits.** In 2019, annual premiums ranged from approximately $600 to $2,400 for insured monthly benefits between $9000 and $24,000 (Ökotest, 2014). The average insured annual benefit was €13,296, with average annual premiums of €924 (Morgen & Morgen, 2021). Because private and public DI benefits are not means-tested, they do *not* crowd each other out; rather, private ODI supplements public WDI or, before the reform, public ODI coverage. The industry relies on its own medical examiners to assess applicants (BBP, 2020).

**Market Structure.** The market is characterized by contractual freedom. Numerous online platforms offer personalized advice, allowing extensive customization of policy components. Consequently, even for identical applicants and coverage conditions, premiums vary widely across insurers (Ökotest, 2014). In 2002, only 26 private ODI insurers operated in Germany, with the three largest (Allianz, Debeka, and ERGO) accounting for roughly two-thirds of new policies sold (Morgen & Morgen, 2021). This concentration suggests a monopolistic market structure, consistent with our model assumptions. Evidence on industry profits supports this interpretation, though detailed firm-level figures are proprietary and a well-kept secret. BaFin, the German Financial Supervisory Authority, reported that in 2000 the aggregate profit share for the life insurance segment (of which ODI is a major component) amounted to 33% of total premiums (BaFin, 2004). By regulation, policyholders must receive at least 75% of this surplus, leaving slightly more than 8% of premiums as retained surplus for insurers. Within the 33%, the largest components were 'capital returns' (20%) and 'risk returns' (7%), the latter reflecting underwriting gains from lower-than-expected claims and lapse rates (BaFin, 2002).

## 3 Impact of the 2001 Reform on Public DI Inflows

This section provides evidence on the effectiveness of the 2001 reform by estimating its impact on inflows to public disability insurance (DI).

**Data.** We use administrative data from the SPI covering the years 1995–2019, disaggregated by year, region, gender, and birth cohort. The outcome variable is *total public DI inflows*, the sum of WDI and ODI, as separate data are unavailable. We normalize inflows by cohort size using population data from the *German Federal Statistical Office*, focusing on birth cohorts from 1954 to 1966, aged 29–59 during the observation window.

**Difference-in-Differences Method.** We compare cohorts for whom public ODI was abolished (treatment group) with those grandfathered under the old system (control group) using the following DD specification:

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

where $y_{ct}$ denotes the share of new public DI recipients in cohort $c$ and year $t$; $D_c$ is a treatment indicator; $T_t$ is a post-reform dummy equal to one after 2000; $\delta_t$ and $\rho_c$ are year and cohort fixed effects; and $\epsilon_{ct}$ is the error term, clustered at the cohort level. The identification assumption is that, absent the reform, treated and control cohorts would have exhibited parallel trends in DI inflows. Our setting does not feature staggered treatment timing.

**[Insert Figure 2 about here]**

**Results.** Figure 2 presents an event-study version of equation (1), replacing $T_t$ with a full set of year dummies (2000 as baseline). The five pre-treatment years show stable, statistically indistinguishable trends between treated and control cohorts. In contrast, inflows decline sharply beginning in 2001, continuing linearly in subsequent years to point estimates exceeding –0.2 percentage points (ppt), or roughly a 35% decline relative to the pre-reform mean.

Since 2011, the inflow differential has remained stable and highly significant at more than –0.2 ppt, representing the long-run effect of the reform. If the concurrent benefit reductions affected inflows, these estimates would represent lower bounds (Appendix B2).

Appendix Figure C1 shows separate event studies by gender. Pre-reform trends are again flat and indistinguishable, while post-reform inflows decline sharply among treated cohorts. The effect is substantially larger for men, who are more likely to qualify for DI due to stronger labor market attachment and higher exposure to physically demanding occupations.

Appendix Table C1 reports DD estimates from equation (1). Panel A shows the full sample, Panel B men, and Panel C women. Each column corresponds to a separate DD model consistent with the event-study estimates. The results are robust to including cohort and year fixed effects and an indicator for East Germany. On average, the decline in male inflows corresponds to about 20% relative to the control mean, compared with roughly 10% for women. The long-run effects are approximately twice as large (see Figure C1).

**Validation.** Appendix C2 validates these findings using representative microdata from the German Socio-Economic Panel Study (SOEP) and an alternative identification approach. The SOEP measures the *stock* of DI recipients, enabling a Regression Discontinuity (RD) design using birth year as the running variable. The identifying assumption is that no other discontinuous factor around 1961 affected DI receipt. We employ a similar RD approach below to estimate the reform's impact on private ODI take-up using the representative SAVE.

**Financial Impact of a Health Shock.** Appendix Table C4 examines the financial consequences of work-limiting health shocks using SOEP data before and after the reform. These results inform the private ODI demand estimates below. As expected, work limitations strongly predict subsequent public DI take-up, reduced employment, lower well-being, and a decline in total annual pre-tax household income of about €4,117, consistent with Di Meo and Eryilmaz (2024). The interacted coefficients with the "treated" indicator are sizable but imprecise for DI receipt, employment, and well-being, while the income effect equals only 0.4% of the mean. Thus, these estimates provide little evidence that the reform materially altered the household-level financial impact of health shocks, given Germany's co-existing social insurance programs and intra-household risk sharing.

# 4 Impact of the 2001 Reform on the Private Individual ODI Market

This section studies how the 2001 reform interacted with the private ODI market in a reduced-form setting. We also provide several stylized facts on private ODI take-up in Germany, one of the world's largest individual markets, in the context of its regulation; see also Section 2. Building on Braun et al. (2019), Section 5 then uses a structural equilibrium model to capture the joint interactions between public WDI, ODI, the safety net, and the reform.

## 4.1 Impact of the 2001 Reform on the Private Individual ODI Market

We first test whether treated cohorts purchased private ODI at higher rates than control cohorts to compensate for lost public ODI coverage. Pre-reform, public ODI may have crowded out private ODI demand, consistent with long-standing evidence of public programs crowding out private coverage (Cutler and Gruber, 1996; Clemens, 2015).

**Data.** We use representative survey data from SAVE (Coppola and Lamla, 2013). It contains rich information on socio-demographics, preferences, savings, retirement, and health, variables typically unobserved by researchers and insurers. Using SAVE, we (a) mimic the risk classification used by private ODI insurers, (b) assess the role of private information in take-up, and (c) elicit take-up parameters by income and health for model calibration.

**Sample Selection.** We use all SAVE waves reporting private ODI coverage from 2001–2010 (except for 2002 and 2004), focusing on employees under age 60. We exclude civil servants and the self-employed, who were unaffected by the reform. Table D1 summarizes the sample: 32% of households hold private ODI policies, the average age is 45, and 33% have 13 years of schooling ("high school degree"). Cohorts are identified by birth year.

Using the RD approach from Appendix C2, Figure 3 shows how the reform affected private ODI take-up. Appendix Figure D2 presents robustness checks for narrower samples: SPI contributors, childless households, and one-person households. The x-axis plots birth year; the y-axis, private ODI coverage. Unconditional scatters are overlaid with linear fits on both sides of the cutoff; polynomial robustness checks are shown in the Appendix.

**[Insert Figure 3 and Table 1 about here]**

The figures reveal two main findings. First, private ODI take-up increases with birth year, implying that younger cohorts are more likely to be insured. This pattern reflects both reforms in the 1980s–1990s that reduced public social insurance generosity and extensive public campaigns encouraging private retirement and disability coverage (Burkhauser et al., 2016). Younger applicants are also healthier and thus less likely to face risk surcharges, pre-existing condition clauses, or outright denials (Section 2). Second, there is no clear discontinuous jump at the cutoff for treated cohorts.

As shown in Table 1, local polynomial RD estimates indicate modest and mostly imprecise effects. The full-sample and SPI-insured models yield positive estimates of roughly 14 ppt (Columns (1)-(2)), while the estimates for childless households are near zero and even negative. Estimates for one-person households are around 25 ppt and marginally significant. These results are consistent with Seibold et al. (2025). Relative to the 21% control-group take-up rate, the 95% confidence interval indicates that population-level coverage is unlikely to exceed 55% due to the reform.

Robustness checks vary bandwidths (Figure D3), test for discontinuities in covariates (Figure D4), inspect the running-variable density (Figure D5), and adjust for discretization and polynomial order (Figure D6).

## 4.2 Stylized Facts on the German Individual Private ODI Market

An estimated 14 ppt increase in take-up is modest, leaving most employees' occupational disability risk uninsured. Why? Because the observed equilibrium outcome reflects both demand and supply under existing regulation. Thus, section 5 uses a structural equilibrium model to disentangle these forces. We first summarize key representative stylized facts from SAVE and SOEP data. These parameters are later fed into the model.

**Take-Up Predictors.** Appendix Table D2 reports socio-demographic means for private ODI holders by treatment status. Consistent with market structure, those purchasing private ODI are more likely to be married, have children, live in larger households, be full-time employed, better educated, white-collar, and higher income. Multivariate regressions (Appendix Figure D1) confirm these as significant predictors. Privately insured individuals are also healthier, reporting better self-assessed health and lower incidence of chronic conditions such as stroke, cancer, or hypertension, and are less likely to smoke. These patterns are consistent with risk rating and selective coverage denials. ODI holders also report stronger precautionary savings motives.

**Health Risk Score.** SAVE records multiple health measures: self-assessed health, health satisfaction (0–10 scale), and major conditions (heart disease, stroke, cancer, hypertension, hyperlipidemia, chronic lung disease), along with smoking, doctor visits, and hospital nights. These correspond closely to insurers' health assessment forms (Appendix Figure B2). To condense these measures into a single metric, we construct a health risk score via principal component analysis (Jolliffe, 2002), following Ghili et al. (2024) and Atal et al. (2025). The normalized score (0–1) mimics insurers' risk rating (Appendix Figure D7).[5] The score distribution is left-skewed with a long right tail, consistent with Karlsson et al. (2024) and Handel et al. (2024).

**ODI Take-Up Gradients.** Figure 4 summarizes key take-up patterns. The y-axis shows ODI take-up. The x-axis shows health risk quintiles from the health risk score; separate lines denote income quintiles. In Section 5, the model calibrations will target these 25 take-up parameters. As seen, take-up declines with worse health, as expected under risk rating and coverage denials. Across all health levels, take-up rises with income: the lowest income quintile exhibits rates between 10 and 25%, while the richest quintiles exceed 40 to 50%. Thus, both income and health gradients are pronounced.

One follow-up hypothesis, left for future research, is whether the sick and poor—who are routinely denied coverage by private ODI insurers—have instead purchased substitute insurance products without risk rating or coverage denial. While we are unable to test this hypothesis using microdata directly, we obtain a time series of all private insurance policies sold by year from the *German Insurance Association* (GDV; see Appendix Figure E5). As seen, the first announcement of the reform in 1998 coincided with spikes in the sales of endowed life insurance and deferred annuity policies. We consider this suggestive evidence in support of the hypothesis.

**[Insert Figures 4 and 5 about here]**

**Lifecycle Risk of Work Incapacity.** Using the long-running SOEP panel, we follow individuals observed working full-time between ages 25–35 and again between 55–60 to estimate the *lifecycle risk* of work incapacity, following Burkhauser and Schroeder (2007). Figure E4a plots the probability of ever reporting a severe health limitation by quintiles of health satisfaction and

[5] Health measures are not available for all years; see Table D1. To maximize the sample size when computing empirical moments for model calibration, we include self-employed individuals and those aged 60 to 62 here. The earliest retirement age increased gradually to 63 from 2004.

household income. The risk is substantial—49% for the least healthy, 8% for the healthiest—and decreases sharply with income, from 31% in the poorest quintile to 10% in the richest. Figure E4b adjusts for socio-demographics, education, and occupation (but not income or health), flattening the curves but preserving a strong income gradient. Lifecycle risks exceed 20% for most groups, consistent with Meyer and Mok (2019) for the U.S.

**Private Information.** Finally, we test for evidence of private information in take-up decisions. Figure 5 plots ODI take-up by health risk quintile, differentiating by whether respondents *expect to retire and receive SPI benefits before age 60*, a proxy for expected work disability, since disability is the only SPI benefit available before 60. Across the entire health distribution, those expecting early retirement are more likely to hold ODI. While the positive correlation test requires claims data (Cohen and Siegelman, 2010), the pattern is consistent with adverse selection.

The next section takes this suggestive evidence on private information to inform our modeling decisions. The model's in-sample calibration will target the 25 empirical ODI take-up parameters by income and health (Figure 4) as well as the lifecycle work incapacity probabilities (Figure E4b). Finally, for out-of-sample validation, we will first simulate the core reform. Then, we will simulate further policies to decompose the demand- and supply-side drivers of limited public–private interaction.

## 5 An Equilibrium Model to Study Demand and Supply Responses

This section develops a new version 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). The model very well approximates key elements of the individual private ODI market in Germany for the following reasons:

First, it is crucial to explicitly model both the demand and supply sides to understand the driving forces behind the reform's impact on private ODI take-up at market equilibrium. In particular, coverage denials are an integral part of this nongroup insurance market, as regulation does not guarantee issue of policies. As an alternative to explicitly modeling supply and demand, sufficient statistics approaches are well-suited for marginal policy changes but less so for fundamental reforms like this one (Kleven, 2020).

Second, on the supply side, the model endogenizes the ODI policy design (the share of lost income insured, loadings, and profits). As the reform entirely abolished public ODI, endogenizing supply-side responses in the co-existing private market appears essential. Moreover, modeling ODI policies is relatively straightforward, as ODI is a pure financial (long-term) contract resembling life insurance (Fang and Wu, 2020) and German long-term health insurance (Atal et al., 2025), all of which follow very similar private insurance regulations. Provider networks or bargaining do not exist in these markets, where employees insure the risk of occupational disability over their working lives.

Third, on the supply side, endogenizing coverage denials follows a key institutional characteristic of the German private ODI market. As discussed in Section 2.3, data sharing between insurers exists; consequently, most applicants first inquire policy conditions anonymously and informally (”pre-underwriting”). However, even numbers on denials of formal applications indicate that one-sixth of all applicants are either entirely denied coverage or incur pre-existing condition clauses (Morgen & Morgen, 2021). According to industry experts, a large share of formal and informal applicants are rejected each year (Versicherungsbote, 2014), and a major audit study revealed that 81% of applicants would not be offered the desired level of coverage (Ökotest, 2014). Given those numbers and institutional facts, it seems paramount to model coverage denial as a key ingredient of the supply side; by doing so, we relax the key assumption of no coverage denials in Seibold et al. (2025). Moreover, the model below incorporates risk rating for 750 risk cells, which reasonably approximates reality (Isenbart and Münzner, 1994).

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Finally, on the demand side, we model the key interplay between public WDI, public ODI, the means-tested cash transfer program, and private ODI. As discussed, for the most basic, low-income jobs, occupational disability equals work disability (see Figure 1), and thus ODI provides little insurance value, given the existence of public WDI and the means-tested cash transfer program, which provides a safety net and consumption floor. Thus, co-existing social insurance is a key determinant of demand for private ODI. The modeling of the demand side also incorporates private information, motivated by economic theory and empirical evidence of adverse selection (see the previous section).

Appendix E1 provides possible market equilibria and optimal insurance policies.

## 5.1 Quantitative Model

### 5.1.1 Individual's problem

In reality, individuals purchase private ODI insurance at an average age of 32 and keep it until age 64; occupational work disability occurs at an average age of 46 (Morgen & Morgen, 2021). Accordingly, we simplify an individual's decision-making problem into two main time periods, see Figure 6.

**[Insert Figure 6 about here]**

**Period 1: Labor Market Entry, Endowments, Disability Risk.** In Period 1, between 25 and 35, individuals enter the labor market, draw a health endowment ($h$), an occupation ($o$), and a wage ($w_1$), which are jointly distributed with density $f(h, w_1, o)$. In Period 1, individuals decide on how much to consume ($c_1$) and save for the future ($s$): $c_1 = w_1 - s$.[6] Further, individuals' disability risk *type*, $\theta^i_{h,w,o}$, $i = b, t$, with $\theta^b < \theta^t$, is revealed to the individual, but not the insurer who only observes $h$, $w$ and $o$. However, these observables—used by insurers to calculate risk cells and risk-rated premiums—are only noisy indicators of the actual disability risk, $\theta^i_{h,w,o}$, which remains private information. With probability $\rho = b$(ottom), individuals realize a low risk of disability, and with $1 - \rho = t$(op), a high risk.

**Private ODI Take-Up.** Then, individuals apply for private ODI, and decide whether to purchase private ODI in case of an offer.[7] The insurer calculates optimal policies (Appendix E1),

[6] We assume that everyone survives Period 1.

[7] Note that the model abstains from insurer learning through selective applications (Lucarelli and Saltzman, Lucarelli and Saltzman), that is, it assumes that all individuals apply. We thank an attentive referee for this remark.

conditional on $h, w, o$, and either rejects applicants or offers private ODI contracts $\{\Pi(h, w, o), b\}$, where $\Pi(.)$ is the insurance premium, and $b$ are contracted benefits. For each risk group, the insurer designs one policy for low-risk types and one policy for high-risk types, without knowing with certainty the actual risk type of each applicant. The contract design follows participation and incentive constraints; see below.

**Modeling Social Insurance Interaction.** Individuals make private ODI take-up decisions, anticipating that, in Period 2—during their main work lives—health and income shocks may hit them. A health shock may lead to full work disability or occupational disability. Following the institutional setting, work disability implies that individuals cannot earn *any* labor income but qualify for public WDI, the basic work disability scheme that existed pre- and post-reform and pays about 30% of the wage as a public benefit; see Table B1. Occupational disability, by contrast, implies work incapacity in the *previous occupation* or part-time work capacity. We model occupational disability as individuals earning half their previous wage. In addition to health shocks, individuals know that (health-independent) income shocks may hit them, e.g., through a recession. Income shocks have density $q(\kappa)$, with $\kappa \in [\underline{\kappa}; \bar{\kappa}]$ and create additional uncertainty over the lifecycle. Depending on the size of the income shock and their income position, individuals may quality for the means-tested cash transfer program. In summary, when making private ODI take-up decisions, individuals consider possible health and income shocks that may lead to eligibility for public WDI or the means-tested cash transfer program.

**Period 2: Work Lives and Realization of Shocks.** Period 2 represents the main work lives of individuals and ranges roughly from age 35 until the earliest possible retirement date. In this period, individuals face the risks of income shocks (-$\kappa w$). Further, work-disabling health shocks materialize. With probability $\rho$=b, the individual remains healthy and earns $w_2$. With probability $(1 - \rho)$, a health shock results in work incapacity, parameterized by a loss function $l(h)$ that operates on $w_2$, as detailed below. In case of full work disability, if individuals purchased private ODI in Period 1, they receive public WDI *and* the private ODI benefit ($b$). Following reality, the latter tops up the former, and does not crowd it out. In case of occupational disability, individuals earn $0.5w_2$ and also receive private ODI benefits if they purchased private ODI.

Following reality, if individuals' total income falls below the consumption floor, they receive a top-up from the means-tested cash transfer program, $\Psi$, up to the consumption floor. This is true, regardless of whether the individual remains healthy but experiences an income shock or incurs a health shock that may lead to full work disability or occupational disability. Moreover, individuals always enjoy utility from appreciated Period 1 savings $(1+r)s$. If they purchase private ODI, they must pay the risk-rated insurance premium, $\Pi$, and self-insure otherwise.

**Decision Problem.** Individuals thus solve the following utility maximization problem, where we omit subscripts for readability:

$$U(h,w,o) = \max_{s,c_{DI},c_0} u_1(c_1) + \beta[\rho u_2(h,w_2,o,\theta^b,\Pi,b) + (1-\rho)u_2(h,w_2,o,\theta^t,\Pi,b)] \tag{2}$$

with

$$u_2(h,w_2,o,\theta^i,\Pi,b) = \int_{\underline{\kappa}}^{\bar{\kappa}} \{u(\kappa w_1) + \alpha[\theta^i u(c_{DI}) + (1-\theta^i)u(c_0)]\}q(\kappa)d\kappa$$

subject to

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

$$c_{DI} + \kappa w_1 = w_2 - l(h)w_2 + (1+r)s - \Pi + b + \Psi_{DI} \tag{4}$$

$$c_0 + \kappa w_1 = w_2 + (1+r)s - \Pi + \Psi_0 \tag{5}$$

where $r$ is the real interest rate and $\alpha$, $\beta$ discount factors. $c_{DI}$ is consumption under work incapacity with private ODI $\{\Pi, b\}$, while $c_0$ is consumption when healthy. $\Psi$ is the cash transfer via the means-tested program, with $\Psi_{DI} = max[0, C - (-\kappa w_1 + w_2 - l(h)w_2 + (1+r)s - \Pi + b)]$ under disability and $\Psi_0 = max[0, C - (-\kappa w_1 + w_2 + (1+r)s - \Pi)]$ when healthy.

### 5.1.2 Insurer's problem

Applicants for private ODI have to indicate $h$, $w$, and $o$ on their application, but the true disability risk $\theta^i_{h,w,o}$ remains private information. The insurer operates in a monopolistic market and maximizes profits $\Xi$ as follows:

$$\Xi(h,w,o) = \max_{\Pi,b} \rho[\Pi^b - \theta^b[\lambda b^b + kI(b^b > 0))] + (1-\rho)[\Pi^t - \theta^t(\lambda b^t + kI(b^t > 0))] \tag{6}$$

where the insurer's variable costs are $\lambda$ (e.g., claims processing) and the insurer's fixed costs are $k$ (e.g., broker commissions). The insurer either denies coverage or offers optimal contracts $\{\Pi(h,w,o), b\}$ to profitable applicants. As mentioned, if profitable, the insurer offers two optimal contracts for each risk cell, one for high and one for low risk types, given the incentive compatibility constraint (*"I prefer my policy over the policy designed for the other risk type"*)

$$u_2(s,\theta^i,\Pi^i,b^i) \geq u_2(s,\theta^i,\Pi^j,b^j) \ \forall i,j \in \{t,b\}, i \neq j \tag{7}$$

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

$$u_2(s, \theta^i, \Pi^i, b^i) \geq u_2(s, \theta^i, 0, 0) \ \forall i \in \{t, b\} \tag{8}$$

### 5.2 Parameterization of Model

The model's parametrization proceeds in two steps, as summarized in Table 2. First, we assign parameter values directly from the data, without solving for equilibria. Second, we calibrate the remaining parameters by minimizing the distance between the empirical and model-generated moments. Following Braun et al. (2019), we do not formally estimate the model, as solving it is computationally intensive. In particular, the menus of optimal insurance contracts must be computed for 750 distinct risk groups, defined over $(h, w, o)$, given risk-rated premiums. These risk cells provide a realistic degree of heterogeneity and approximate the observed market structure, see Section 2.

We employ a standard utility function with constant-relative risk aversion $u(c) = \frac{c^{1-\sigma}}{1-\sigma}$. We set the risk aversion parameter $\sigma$ to 2 and the interest rate to zero, following Braun et al. (2019). The discount factor $\beta$ is set to 0.94. As fixed ($k$) and variable ($\lambda$) administrative costs, we take industry averages of 3% of lifetime and 10% of annual premiums, respectively; see Table 2 (Finanzberatung Bierl, 2023).

**[Insert Table 2 about here]**

**Health Risk.** Figure D7 shows the actual health risk distribution computed from representative SAVE data, see Section 4.2. A beta distribution with $\beta(1.23; 6.83)$ approximates this skewed distribution reasonably well. The mean risk scores by the five income quintiles are in Table E1. We assume that their joint distribution follows a Gaussian copula with parameter $\varphi = -0.29$, chosen to match the data points in Table E1. As shown in Appendix Table E1, our method accurately reproduces the empirical distribution.

**Initial Income.** We extract the wage distribution of employees between the ages of 25 and 35 from the representative SOEP. We then model it as a log-normal distribution, normalize the mean to one, and calibrate the standard deviation to match the Gini in the year 2000.

**Income Uncertainty.** Again, using representative SOEP data, we calibrate the distribution of income uncertainty for Period 2. We then calculate the post-government household income between 45 and 55 (middle of Period 2) as a share of net income in ages 25 to 35 (Period 1).

This represents the change in net income over the work lives, driven by changes in household income and household composition. Further, we use the empirical 5% and 95% bounds of that distribution as bounds of a 1-$\kappa$ truncated log-normal distribution in the model. Most individuals' household income increases over their work lives, between Period 1 and 2. [8]

**Work Incapacity Costs.** Under full work disability, individuals cannot earn any labor income but receive public WDI benefits, roughly 30% of their previous income (Table B1). Occupational disability, by contrast, implies that individuals lose 50% of their previous wage, see Section 2. We thus parameterize the costs of a health shock in the form of a loss function:

$$l(h_i) = P(WDI|WDI \vee ODI) * 0.7 * w_2 + P(ODI|WDI \vee ODI) * 0.5 * w_2 \tag{9}$$

**Work Incapacity Risk.** The probability of a work-disabling health shock differs by occupation and health. Suppose employees in the most basic (low-wage) occupation, Category 1 in Figure 1, experience a work-limiting health shock. Following Figure 1, it will always lead to WDI, as there is no wedge between a higher occupational status and the ability to perform the most basic work in the economy. By contrast, for the most skilled and highest-status occupations, Category 6 in Figure 1, we assume that employees' occupational disability risk would be 80%, (and the full work disability risk 20%), given a health shock. We then extrapolate these conditional ODI risks for the intermediate categories. Thus, our assumed work vs. occupational disability risks by quintiles 1 to 5 are $P(WDI|WDI \vee ODI) = [0.2, 0.4, 0.6, 0.8, 1]$. Inserting those probabilities in equation (9), we obtain the following expected disability costs as a share of previous wages by risk quintiles: $l(h_i) = [0.54, 0.58, 0.62, 0.66, 0.7]$. This is our parameterization of the key criticism of public ODI by the *Federal Audit Office* and all democratic parties at the time, which led to its elimination (Tempel, 2018): Namely, that all employees fund public ODI via payroll taxes[9], but higher-status employees benefited disproportionately.

**Matching Simulated Moments** In the second step, we solve for model equilibria and choose key parameters to minimize the distance between empirical and model-implied moments. The calibration targets two sets of moments. The first, based on the representative SOEP data, captures 25 lifecycle work-incapacity probabilities by income and health-risk quintiles (see Figure E4b). The second, based on the representative SAVE data, captures 25 private ODI

[8] Negative numbers for $\kappa$ represent increases in income.

[9] As the payroll tax has a contribution ceiling, the funding is actually regressive too

take-up rates by health and income quantiles (see Figure 4).

We employ a two-step algorithm to minimize the distance between the empirical and model-implied private ODI take-up rates. To do so, it searches for the optimal combination of the type-specific probabilities of lifecycle work incapacity and the fraction of good types, $\psi$, where the type-specific probabilities of work incapacity constitute private information. The calibration aims to find the combination of $\psi$ and type-specific work incapacity that yields a close model-based match between ODI take-up rates by health and income and the respective data moments (Figure 4), while also matching the population-based work-incapacity parameters by income and health from the data (Figure E4b). Appendix F provides additional details.

**[Insert Figure 7 about here]**

The calibration yields a share of good types of $\psi = 0.62$. As shown in Figure 7, we achieve a strong in-sample fit: model-generated ODI take-up rates by income and health closely match their empirical counterparts.

#### 5.2.1 Out-Of-Sample Model Validation and Reform Simulation

The previous subsection demonstrated a strong in-sample model fit for the calibrated model parameters. This subsection provides an out-of-sample validation with untargeted parameters. Specifically, we use the model to simulate the 2001 reform and compare the take-up effect to the estimated reduced-form effects reported in Section 4.1. As no pre-reform SAVE data exist, we simulate the reverse reform effect using post-reform data and then simulate pre-reform take-up. We then benchmark simulated take-up with the empirical causal reform estimates. To specify the pre-reform costs of an occupational disability, we use individuals' expected, self-reported pension replacement rates from the representative SAVE (12% to 18%), see Appendix B2, which also contains a stylized benefit simulation that produces very similar replacement rates.

Figure 8 displays an average simulated pre-reform private ODI take-up rate of 19% and a post-reform rate of 33%, implying a 14 ppt increase attributable to the reform. We benchmark this simulated reform effect against three independent out-of-sample estimates: First, the simulated effect closely matches the reduced-form point estimate from Figure 3 and Table 1, despite the latter's statistical noisiness. Second, in the representative SAVE data, outside of any estimation, 22% of those in the grandfathered (control) cohorts hold private ODI, compared with 42% among those (treated), born after 1960. Third, using rating agency data, Seibold et al. (2025) document that the private ODI take-up rate rose from below 10% in 1997 to 26% by 2015.

Taken together, these sources suggest that the model's simulated reform effect is broadly consistent with the observed empirical patterns and closely aligns with external data and the representative reduced-form estimates.

#### 5.2.2 Policy Simulations

Finally, we use the model to evaluate policy counterfactuals. In particular, we combine the actual reform with select complementary policies and simulate their implied take-up effects. Specifically, we assess how the reform would have affected private ODI take-up rates, both in the overall population and by health categories, had policymakers implemented it *jointly* with:

***Demand-Side Policies:***

- An increase in WDI benefits by 10%.
- An increase in WDI benefits by 25%.
- A reduction in means-tested cash transfer benefits by 10%.

***Supply-Side Policies:***

- Policies to reduce variable administrative costs, e.g., through minimum benefit ratios.
- Policies to reduce fixed administrative costs, e.g., through banning broker commissions.
- Policies to reduce private information, e.g., through genetic testing.

Simulating these hypothetical policies provides insights into both the direction and magnitude of the equilibrium take-up response. It also helps clarify the model's internal mechanisms and highlights the relative importance of supply- and demand-side forces, such as the interaction between overlapping social insurance schemes and supply-side regulations.

The first three policies vary the generosity of co-existing social insurance strands. They are straightforward to implement and regularly discussed in Germany; variants have already been implemented, see Appendix B. These policies target the demand side.

The following three policies target the supply side and constitute indirect measures. Clearly, eliminating all variable and fixed administrative costs is unrealistic. Moreover, policies that ban broker commissions or impose minimum benefit ratios, similar to the ACA's minimum loss ratios (MLRs), are controversial and may lead to unintended consequences such as cost-shifting or the relocation of business activities (Inderst and Ottaviani, 2012; Braegelmann and Schiller, 2025). However, to our knowledge, there is no systematic empirical causal evidence that such unintended consequences have indeed occurred (Karaca-Mandic et al., 2015; Cicala et al., 2019;

Born et al., 2023; Kaiser Family Foundation, 2024). At the same time, the rise and growing importance of online-only and direct-to-consumer digital insurers are likely to reduce broker fees and administrative costs across the industry over time. The simulation exercise, therefore, also captures the gradual impact of these new products on take-up.

This discussion also extends to private information, which, of course, can never be entirely eradicated. Nevertheless, the relevance of private information for the long-term take-up equilibrium appears paramount. We therefore simulate the effect of eliminating private information on take-up. As one (in our view, unethical) potential policy example for reducing information asymmetries, we refer to genetic testing, which has been applied in insurance markets around the world. In Germany, the use of genetic tests or relatives' medical histories has only been prohibited since 2010 under §18 of the Gendiagnostikgesetz (GenDG), with explicit exceptions *permitted* when the insured benefit exceeds €30,000 per year in private ODI.

Figure 8 reports the results of the policy simulation.

**[Insert Figure 8 about here]**

**Expanding Public WDI.** Private ODI take-up would have been one percentage point higher had policymakers simultaneously increased WDI benefits by 10%, and eight percentage points higher under a 25% increase. The mechanism operates through the enhanced value of ODI contracts: higher WDI benefits reduce the likelihood that the means-tested consumption floor crowds out private ODI payouts, particularly among low-income households. Note that expanding public WDI is a very plausible and broadly supported policy alternative.

**Reducing Means-Tested Cash Transfers.** Lowering the consumption floor by 10% would increase private ODI take-up by approximately nine ppt, primarily by enhancing its value for low-income households. As an extreme thought experiment, a complete abolition of this welfare program—though clearly infeasible under current constitutional constraints—would more than triple overall private ODI coverage to 81% (details available upon request). This counterfactual highlights the pivotal role of the public safety net in shaping private ODI demand. The finding stands in sharp contrast to Seibold et al. (2025), who abstract from co-existing social insurance schemes, with direct implications for welfare analysis.[10] Our results indicate that public WDI and means-tested cash transfers act as substitutes for private ODI.

[10]Their Appendix Table A5, column (5) of Panel B, shows a simple DD regression with treated×2005 on 480 cohort-year observations to test whether the 2005 welfare reform affected take-up. From the imprecise estimate whose upper 95% CI includes 33% of the mean, they conclude that "*[...] the generosity of social assistance is not a major driver of private DI take-up.*"

**Reducing Administrative Costs.** The next two bars in Figure 8 simulate a 50% reduction in fixed and variable administrative costs, respectively. Limiting broker fees, a real-world policy with this objective, is currently under active debate, for instance in Germany (BaFin, 2018) and at the European Union level (Reuters, 2023), and has already been implemented in countries such as the Netherlands, the United Kingdom, and Australia (RGA, 2014). According to our simulations, take-up would have increased by an additional six percentage points under a 50% reduction in fixed costs and by seven percentage points under a 50% reduction in variable costs. The potential increase in take-up from lowering administrative costs is bounded at 20 ppt in the unrealistic scenario of zero administrative costs (details available upon request).

**Eliminating Private Information.** Another somewhat unrealistic upper-bound scenario is represented by the last bar in Figure 8. It indicates a potential 20 ppt increase in take-up under the assumption of no private information. In the absence of private information, insurers can tailor type-specific policies, thereby reducing the likelihood of erroneously denying coverage or offering policies optimized for high-risk individuals to low-risk ones. A controversial policy that could, in principle, partially achieve this objective would allow insurers to perform genetic tests for risk rating, as remains the case in the U.S., whereas France and Canada maintain strict bans. Switzerland and the Netherlands have adopted frameworks similar to Germany's, permitting genetic testing above certain insured thresholds.

**[Insert Table 3 about here]**

**Private ODI Policy Design.** Table 3 details how the reform and additional simulated policies would have endogenously affected policy *design*. Analyzing such general equilibrium outcomes on the supply side is a key strength of our model. For instance, increasing public WDI benefits by 25% raises overall take-up by an additional eight percentage points. At the same time, denial rates decrease only marginally, indicating that the rise in take-up is primarily demand-driven. A similar pattern emerges for the simulated reduction in the consumption floor: denial rates remain virtually unchanged, while the "share of cost covered" by private ODI increases. "Insurer profits" rise slightly when the safety net is reduced and decline modestly under a more generous public WDI program. By construction, though reassuringly consistent with expectations, "safety net transfers" increase under the former policy, while DI transfers decline under the latter.

**Heterogeneity by Risk Type.** Panel B of Table 3 differentiates all endogenous outcomes for good and bad risk types, $\theta^b$ and $\theta^t$. The market is adversely selected by unobserved risk types, with only 3% of good risks but 45% of bad risks purchasing policies in the pre-reform era. After the reform, take-up increases by 10 ppt among good and by 20 ppt among bad risk types. Second, the ”share of costs covered” by private ODI is higher for bad than for good risks. Bad risks receive relatively favorable terms because insurers cannot identify them with certainty and must design contracts under incentive compatibility constraints, which limit profits. Interestingly, while ”the share of costs covered” by ODI policies remains constant at 84% for bad risks, it increases by 11 ppt for good risks after the reform, illustrating that good risks demanded and were offered more generous ODI policies. Third, the loading factor, one minus the ratio of the expected value of benefits to premia, decreases for both risk types as a result of the reform, reducing what insurers extract from consumers. Cutting administrative costs in half would further lower the load for bad risks and strongly increase take-up to around 75%.

**Equity.** Figure E6 (Appendix) reports simulated take-up rates by health quintile. First, overall take-up never exceeds 42%, regardless of risk type or policy scenario (except when entirely eliminating private information). Second, none of the simulated policies fully eliminates the health gradient in take-up, although almost all policies would flatten it. Third, reducing the consumption floor and increasing WDI would shift take-up upward in an almost parallel fashion by roughly 10 ppt across most health quintiles.

**Welfare.** We measure welfare in terms of model-based utility (‘utils’) but also by Consumption-Equivalent Variation (CEV). CEV translates welfare differences into interpretable percentage changes in consumption, answering: by what percentage would we need to adjust individuals’ income in the counterfactual scenario to make them equally well off? This money-metric measure complements our utility-based rankings by providing an economically meaningful scale for welfare comparisons. Appendix G provides details.

We acknowledge that utility levels are not cardinal and cannot be interpreted directly in absolute terms. However, differences in welfare are invariant to monotonic transformations of the utility function. As such, the reported values provide a consistent basis for welfare comparison, but not in money-metric terms. Because CRRA has a risk preference factor of 2, utility is negative, and less negative utility implies higher welfare.

Table 3 shows that, despite rising insurer profits due to the 2001 reform, overall consumer utility (which includes insurer profits) decreased slightly as public ODI was cut. However, this lower utility does not account for potentially lower contribution rates due to lower public spending; see below. As seen in columns (3) to (8), welfare could have been higher had policy-makers combined the reform either with increases in WDI benefits or measures to eliminate either fixed or variable administrative costs.

The CEV results in Table 3 reveal a similar pattern, indicating that the 2001 reform would have required an income compensation of 0.4% to make individuals equally well-off. Compared to the post-reform economy in column (2), increasing WDI benefits by 10% or 25% increases welfare and would have required negative compensation of -0.1% and -0.35%, respectively, whereas reducing the consumption floor by 10% would have required income increases of 0.4% to make individuals equally well off. Notably, reducing administrative costs yields only modest welfare gains (0.05%), while perfect information would even reduce overall welfare by 0.1%.

Considering average pre-reform public ODI benefits and the decrease in inflows by a third per cohort per year (Figure B1), a simple calculation suggests that the reform reduced SPI expenditures by €789 million per year, or about €12 billion after 15 years, when the first affected cohorts had basically aged through their work lives. Further assuming that this decrease in public DI spending would be entirely passed through to taxpayers, contribution rates would be one percentage point lower, resulting in reduced employer and employee payroll deductions of €500 per year for an annual income of €50,000. That is, 1% of annual income. This amount would be sufficient to (over)compensate individuals for potential welfare losses in all policy scenarios above. Finally, note that the reform-induced spending reduction of €789 million per year is about 5% of total WDI spending, as the long-run equilibrium was reached in 2012 (Deutsche Rentenversicherung, 2025a). An alternative to lowering payroll taxes could have been to use these savings to increase WDI benefits by 5%.

# 6 Conclusion

This paper examines how a fundamental reform of the German public disability insurance system interacted with the private disability insurance market. To analyze the reform's effects, we combine reduced-form and structural approaches with administrative and survey data. The 2001 reform eliminated access to public occupational disability insurance (ODI) for cohorts born after 1960. The stated motivation, supported by all major political parties, was that the existing system disproportionately protected higher-income and better-educated workers, effectively constituting *"a privilege of insurees with extraordinary education and in elevated positions"* (Bundestag, 1997).

We first document the reform's effect on public disability insurance (DI) inflows using administrative data and event-study methods. The results indicate a long-run decline of approximately one-third in new DI inflows, which materializes about a decade after implementation. We then study interaction effects with one of the largest private ODI markets worldwide. This market is largely unregulated, with risk-rated premiums, medical underwriting, and no guaranteed issue. Most policyholders purchase private ODI policies around labor market entry and keep those policies for roughly three decades, covering the core working years. Contracts are purely financial with one-sided commitment: policyholders can lapse, but insurers cannot rescind coverage.

Consistent with the literature, our reduced-form estimates reveal at most a modest increase in private ODI take-up among cohorts that lost access to public ODI. Because take-up is an equilibrium outcome, we extend and tailor Braun et al. (2019) to the German institutional setting, incorporating key demand- and supply-side forces.

On the demand side, we model interactions with two public programs that shape private ODI demand: public work disability insurance (WDI), which covers *general* work disability, and the means-tested cash transfer program that provides a universal *consumption floor*. These programs are central to understanding private ODI demand and, in particular, differential *uptake* by income. In addition, private ODI provides "social status protection," yielding higher insurance value for higher-income households.

On the supply side, health-risk rating and routine coverage denials generate pronounced variation in take-up across population groups. Higher-income and healthier individuals face lower premiums and lower denial rates.

We then simulate policy counterfactuals to isolate the drivers of equilibrium take-up. On the demand side, increasing WDI generosity and reducing the consumption floor both raise private ODI take-up (by roughly eight and nine percentage points, respectively). Nonetheless, even under these policies, overall coverage remains below 50% and retains pronounced, though flatter, health gradients.

On the supply side, reducing administrative costs and information asymmetries increases take-up. Lower administrative costs primarily benefit higher-risk individuals by lowering premiums and improving affordability. Reducing private information mainly benefits lower-risk types; effects for higher-risk types are mixed—tighter risk classification can raise their premiums and reduce take-up, even as conditional generosity may rise.

Even under favorable scenarios, however, about one-third of the population remains without occupational disability insurance. Achieving substantially higher coverage, if deemed desirable, would require stronger regulation (e.g., individual mandates or community rating) beyond the incremental policies considered here.

A promising avenue for future research is to understand why the U.S. individual private ODI market is substantially smaller while the group market is much larger (cf. Herbst and Hendren, 2024). Moreover, interactions and substitution across private insurance products warrant further study. We find suggestive evidence of sharp increases in capital endowment life insurance policies around the 2001 reform, pointing to potential substitution between ODI and other financial products. While a complete analysis of such private policy interactions is beyond the scope of this paper, understanding how households reallocate across insurance and savings instruments in response to policy changes is an important direction for future work.

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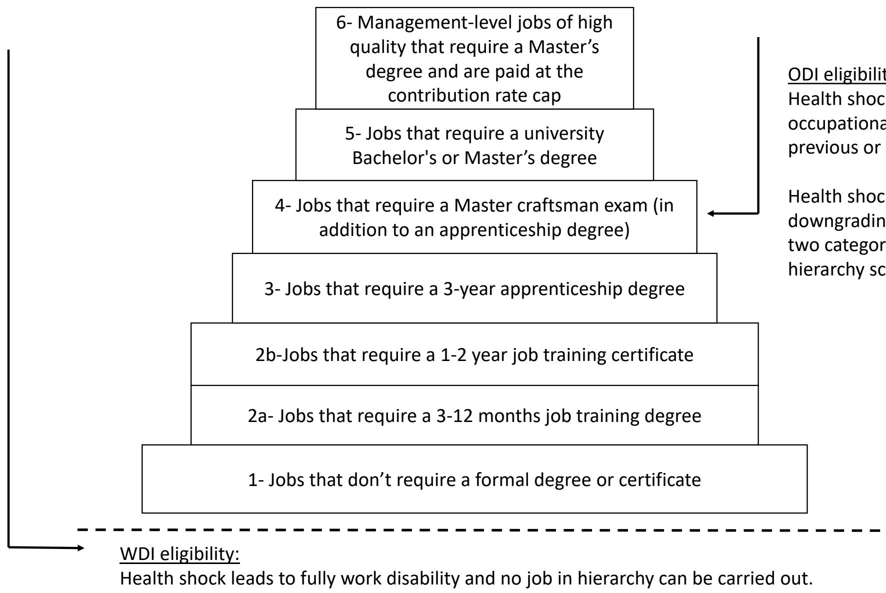
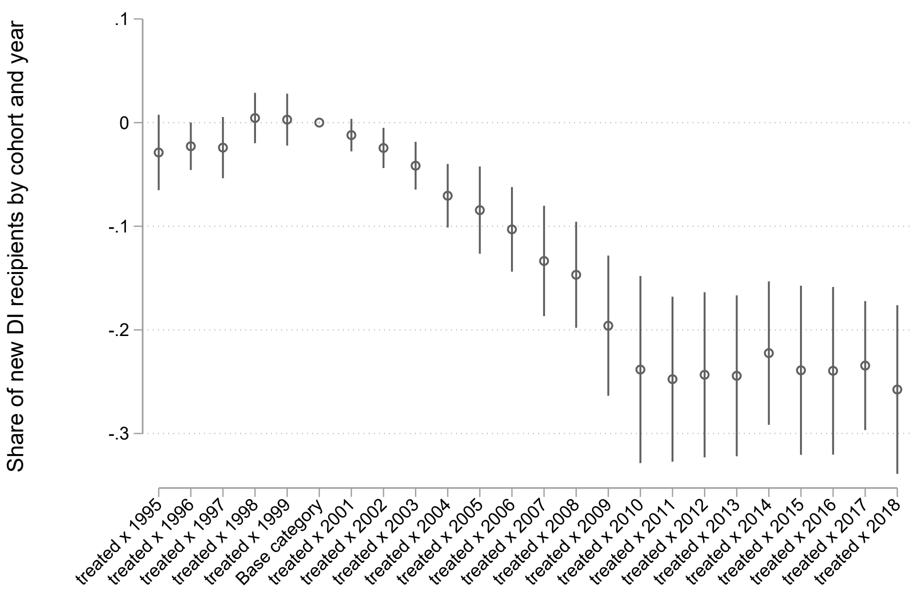
**Figure 1:** Job Hierarchy Scheme Illustrating WDI and ODI Eligibility



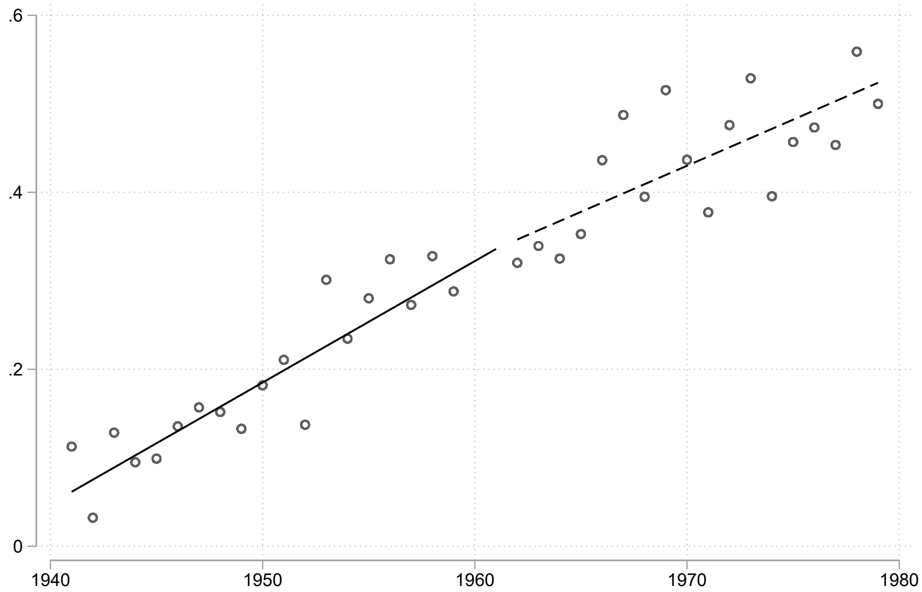

*Source:* own illustration and translation following §240 SGB VI and Deutsche Rentenversicherung (2025b). The Federal Social Court has developed a job-hierarchy scheme under which ODI eligibility is established if a worker is unable to perform their previous job or a job in a lower category due to health issues. For blue-collar jobs, categories five and six do not exist. Appendix Figure B1 provides further details on ODI and WDI edibility.

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



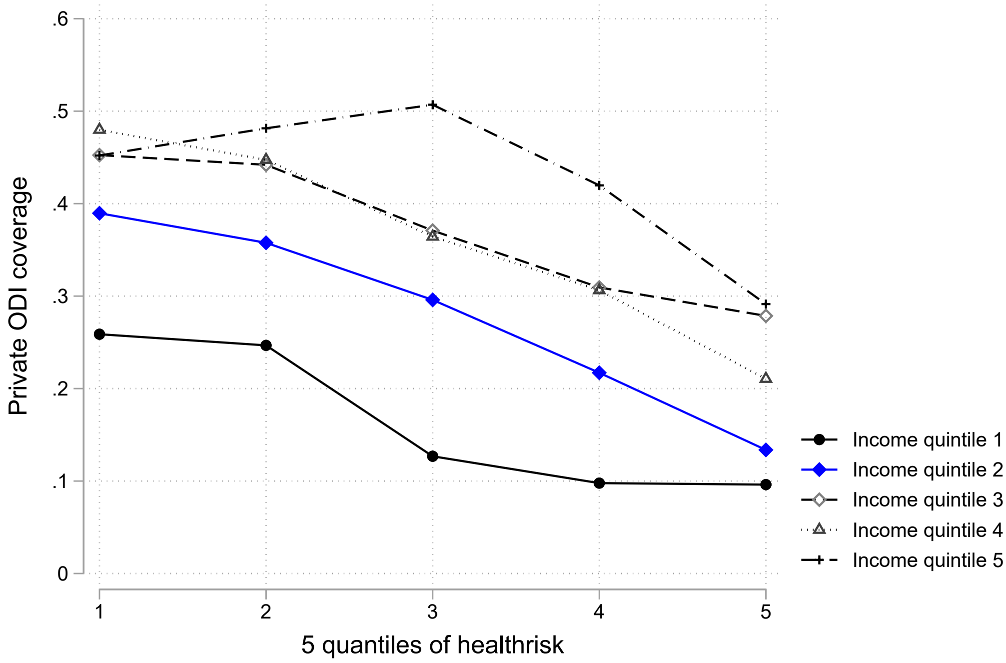

*Source:* Administrative SPI data on new public DI recipients by cohort and year. Treated cohorts are those born after 1960, and the treatment group; grandfathered cohorts are those born before 1961, and the control group. Figure plots $\beta D_c \times T_t$ estimates from equation 1 but with the post-reform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

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



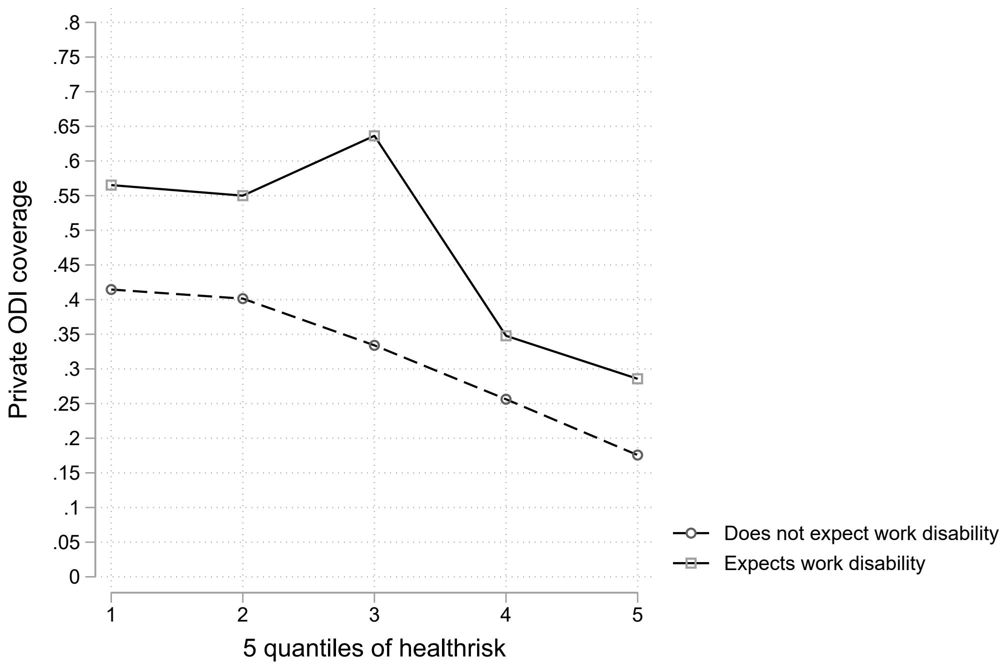

*Source:* SAVE data 2001-2010. The figure shows the raw nonparametric means of private ODI coverage by birth year, overlaid with separate linear trends before and after the cutoff. Other robustness checks vary the sample (Figure D2), vary the bandwidth (Figure D4), study the smoothness of covariates (Figure D5), carry out density plots of the running variable (Figure D6, McCrary 2008), and vary polynomials as well as run donut RDs (Figure D2).

**Figure 4:** Take-Up of Private ODI Policies by Health Risk and Income



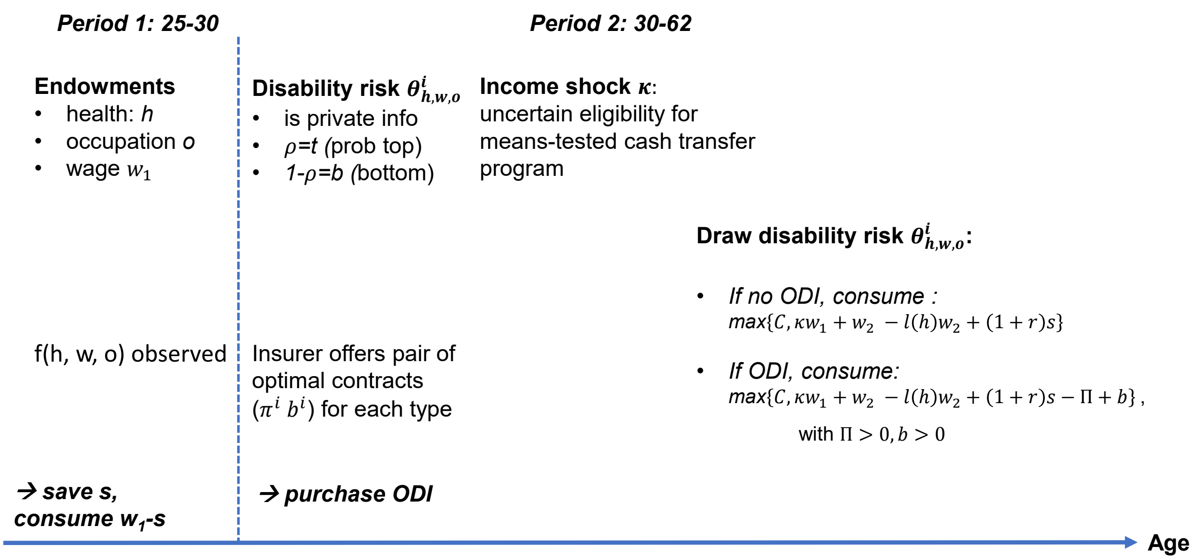

*Source:* SAVE data 2001-2010. Figure plots take-up rates of private ODI policies against the quintiles of the health risk score in Figure D7 and stratifies these curves by the five net household income quintiles.

**Figure 5:** Take-Up by Health Risk and Private Information



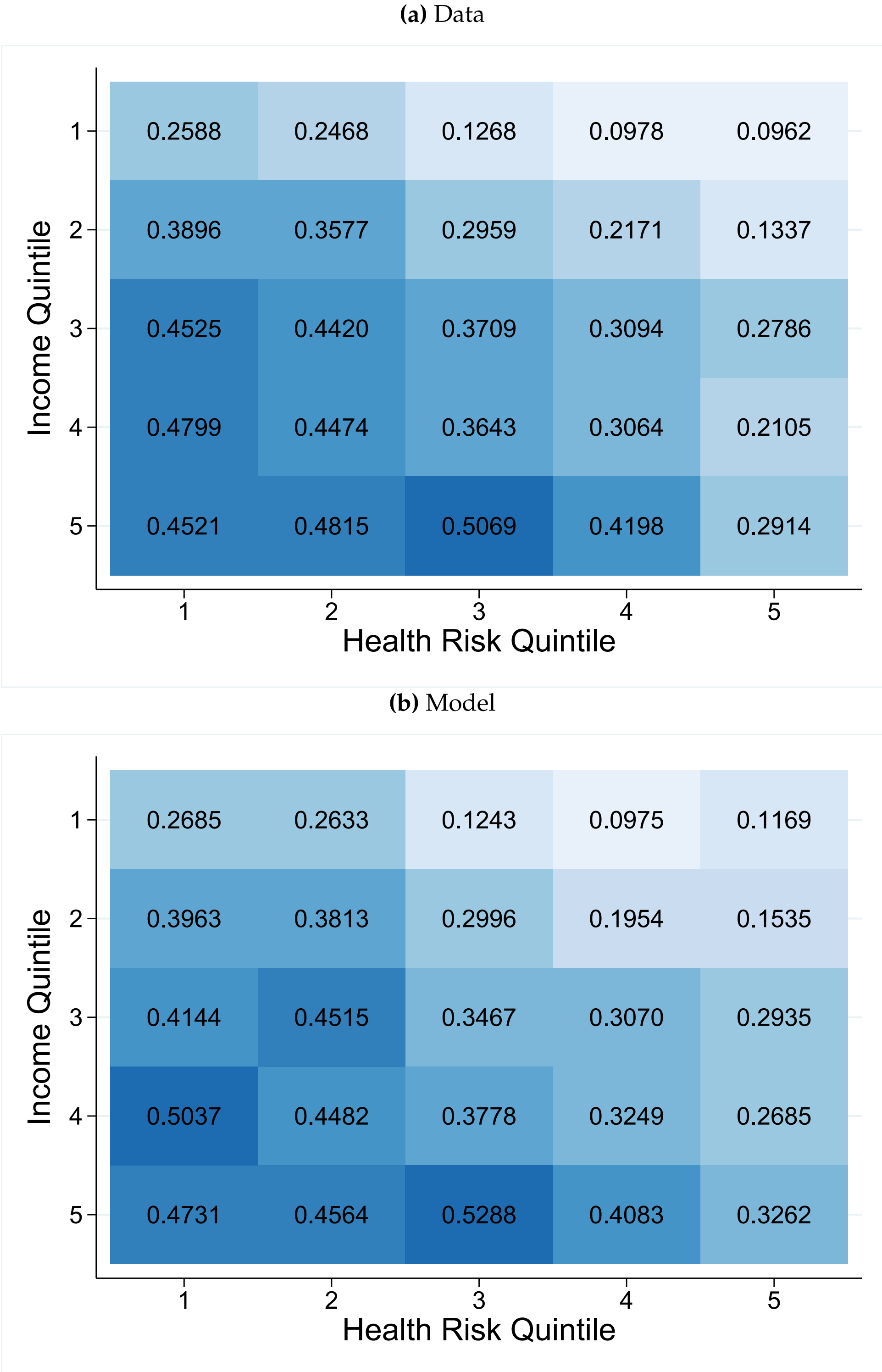

*Source:* SAVE data 2001-2010. Figure plots take-up rates of private ODI policies against the quintiles of the health risk score in Figure D7 and stratifies these curves by expected retirement before age 60. The latter information is directly elicited in the SAVE survey and proxies for expected work disability.

**Figure 6:** Illustration of Lifecycle Time Periods in Baseline Model



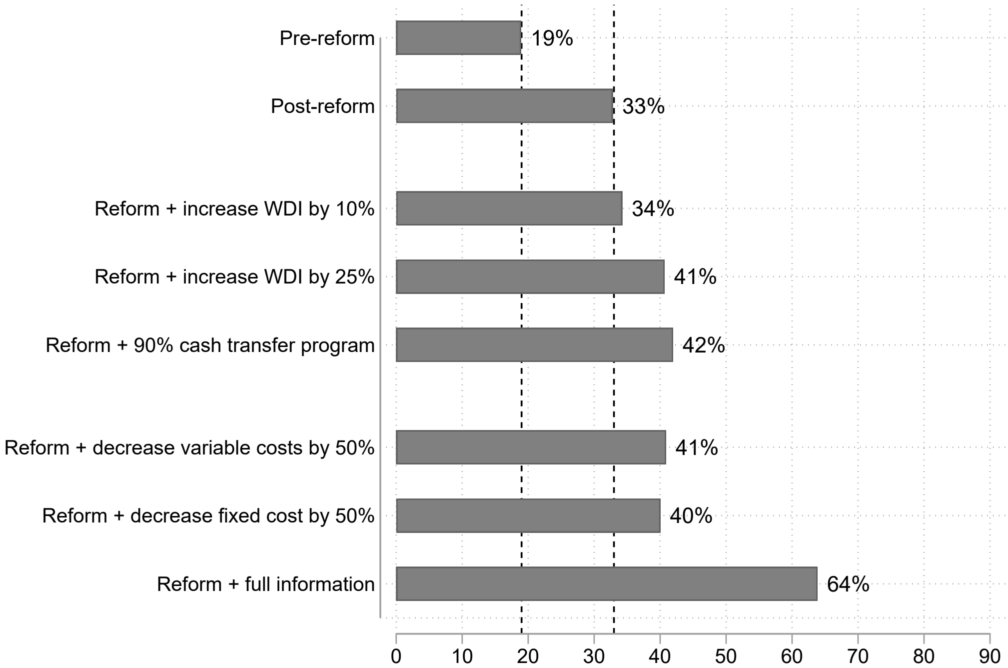

*Source:* The figure illustrates the lifecycle decision-making process of a customized version of the GEM model by Braun et al. (2019). For more details, please see the main text.

**Figure 7:** Model Fit: Private ODI Take-Up by Income and Health Quintiles




*Source:* Figure shows private ODI take-up rates by health risk (columns) and income quintiles (rows). '1' is the healthiest and poorest quintile, whereas '5' is the sickest and richest quintile. Figure (a) shows the raw data from SAVE, and Figure (b) shows the predicted take-up rates by the general equilibrium model.

**Figure 8:** Effect of 2001 Reform on Private ODI Policy Take-Up: Simulation




*Source:* The bars show model predictions for average population-level reform effects. The private ODI take-up is simulated; see the main text for details.

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

| | (1)<br>Full sample | (2)<br>SPI insured | (3)<br>No kids | (4)<br>One-person HH |
|---|---|---|---|---|
| Bias-corrected | 0.140 | 0.141 | -0.018 | 0.246* |
| | (0.0924) | (0.1042) | (0.0372) | (0.1267) |
| Robust | 0.140 | 0.141 | -0.018 | 0.246* |
| | (0.0982) | (0.1150) | (0.0473) | (0.1434) |
| Conventional | 0.108 | 0.129 | -0.035 | 0.202 |
| | (0.0924) | (0.1042) | (0.0372) | (0.1267) |
| Year FE | yes | yes | yes | yes |
| State FE | yes | yes | yes | yes |
| Age + gender | yes | yes | yes | yes |
| Observations | 10,395 | 7,868 | 5,131 | 1,899 |

*Source:* SAVE data 2001-2010. See Coppola and Lamla (2013) for more details about SAVE. Only respondents below 63 are included; civil servants, the self-employed, and birth years 1960 and 1961 are omitted. The tables show point estimates of local polynomial regressions similar to equation (12) (Calonico et al., 2014, 2017, 2018, 2019) using a data-driven optimal bandwidth selection, 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 D3, (Calonico et al., 2020), study discontinuities in covariates (Figure D4), carry out density plots of the running variable (Figure D5, McCrary (2008), and vary polynomials (Figure D6).

**Table 2:** Model Parameters

| | | |
|---|---|---|
| *Parameters set* | | |
| Risk aversion | $\sigma$ | 2 |
| Discount factor | $\beta$ | 0.94 |
| Insurer's variable costs | $\lambda$ | 1.1 |
| Insurer's fixed costs | $k$ | 1.03 |
| Net interest rate | $r$ | 0 |
| | | |
| *Parameters calibrated* | | |
| Health risk distribution | $f$ | $\beta(1.23; 6.83)$ |
| Copula parameter | $\varphi$ | -0.29 |
| Period 1 wage | $w_1$ | $\ln(w_1) \sim (-0.15, 0.55^2)$ |
| Period 2 wage | $w_2$ | $= \gamma w_1, \gamma \in [1.25,\ 1.30]$ (log-spaced) |
| Income shock distribution | $\kappa$ | $\ln(\kappa) \sim \mathcal{N}(-0.34,\ 0.63^2)$, truncated to $[-1.46\ 0.52]$ |
| Means-tested cash transfer | $c$ | 0.12 |
| Fraction good types | $\psi$ | 0.62 |

*Sources:* Set parameters follow Braun et al. (2019) and (Finanzberatung Bierl, 2023). SAVE is used to calibrate the health risk score distribution (Figure D7), and SOEP for the wage distributions, and the income shock. The calibration for $\psi$ is explained in Appendix F. The consumption floor (€ 11,868 p.a., Bundesagentur für Arbeit (2019)) is expressed as the share of permanent income, which is the average gross wage (€ 47,928, Statistisches Bundesamt (2022)) multiplied by the average contract duration (31.5 years). All figures are rounded to two digits for readability.

**Table 3:** Take-Up, Loading, and Risk Insured: Policy Simulations

| | (1) **Pre-Reform** | (2) **Post-Reform** | (3) **↑10% WDI** | (4) **↑25% WDI** | (5) **↓ 10% cash transfer** | (6) **↓ 50% variable admin** | (7) **↓ 50% fixed admin** | (8) **Full Info** |
|---|---|---|---|---|---|---|---|---|
| Panel A: Total | | | | | | | | |
| Take up | 0.1894 | 0.3283 | 0.3427 | 0.4069 | 0.4198 | 0.4092 | 0.4006 | 0.6387 |
| Denial | 0.5452 | 0.3514 | 0.3413 | 0.3258 | 0.3519 | 0.2513 | 0.2540 | 0.2319 |
| Share of cost | 0.8085 | 0.7313 | 0.7114 | 0.6242 | 0.7754 | 0.7295 | 0.7039 | 0.9428 |
| Load | 0.2775 | 0.2890 | 0.2792 | 0.3020 | 0.3126 | 0.2757 | 0.2688 | 0.5095 |
| Profits | 0.0118 | 0.0247 | 0.0230 | 0.0206 | 0.0277 | 0.0291 | 0.0270 | 0.0535 |
| | | | | | | | | |
| Govt total transfer | 0.0726 | 0.0613 | 0.0664 | 0.0740 | 0.0603 | 0.0606 | 0.0608 | 0.0613 |
| DI transfer | 0.0743 | 0.0576 | 0.0634 | 0.0720 | 0.0576 | 0.0576 | 0.0576 | 0.0576 |
| Means-tested transfer | 0.0038 | 0.0037 | 0.0030 | 0.0020 | 0.0027 | 0.0030 | 0.0032 | 0.0037 |
| | | | | | | | | |
| Ex-Ante Utility | -1.4555 | -1.4611 | -1.4597 | -1.4572 | -1.4664 | -1.4602 | -1.4602 | -1.4621 |
| Average CEV (%) | | 0.4394 | -0.1129 | -0.3505 | 0.3551 | -0.0528 | -0.0540 | 0.1081 |
| Panel B: Good risks | | | | | | | | |
| Take up | 0.0283 | 0.1338 | 0.1508 | 0.2446 | 0.2811 | 0.2031 | 0.1909 | 0.6833 |
| Share of cost | 0.2891 | 0.4026 | 0.4021 | 0.3771 | 0.4109 | 0.4336 | 0.4293 | 0.9383 |
| Load | 0.6609 | 0.5855 | 0.6072 | 0.7107 | 0.6607 | 0.6372 | 0.6071 | 0.6153 |
| Panel C: Bad risks | | | | | | | | |
| Take up | 0.4548 | 0.6486 | 0.6587 | 0.6742 | 0.6481 | 0.7487 | 0.7460 | 0.5651 |
| Share of cost | 0.8617 | 0.8430 | 0.8280 | 0.7719 | 1.0359 | 0.8616 | 0.8197 | 0.9519 |
| Load | 0.2382 | 0.1883 | 0.1555 | 0.0577 | 0.0638 | 0.1142 | 0.1261 | 0.2988 |

*Source:* Table shows private ODI take-up rates, the fraction being denied coverage, the share of costs insured, loading factors, insurer profits, consumer utility, transfers and all scenarios by good and bad risk types. All policies in columns (3) to (8) are combined with the 2001 ODI reform. CEV stands for consumption-equivalent variation. See Appendix G for more information. Column (2) reports the percentage income increase required in the post-reform economy for individuals to be equally well-off as in the pre-reform economy (Column (1)). Columns (3) to (8) report the percentage income increase required in the counterfactual economy for individuals to be equally well-off as in the post-reform economy (Column (2)).

# Online Appendix

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# B Further Institutional Details

**Figure B1:** 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.<br><br>5-year waiting period after labor market entry. | Does health status allow 3 hours of work per day in any job? | 100%<br><br>2000: €731<br>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 60 up to 10.3% | Available throughout the entire time period for all cohorts. |
| **Occupational DI (ODI)** | Work disability in last or trained occupation | Social contributions paid in last 3/5 years.<br><br>5-year waiting period after labor market entry. | Does health status allow 6 hours of work per day in previous/trained occupation? | 50%<br>*(same as partial WDI post-2000[1])*<br><br>2000: €584<br>2005: €515 | Same as WDI but is supposed to solely compensate for partial work capacity loss.<br><br>About 12% of gross wage with average age at first receipt of 47, see Appendix B for details. | Cut for cohorts born after 1960.<br><br>Effective insurance value is higher, the higher wage in the last occupation, cf Figure 1. 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, (Deutsche Rentenversicherung, 2023b). See main text for more details. ODI was abolished for cohorts born after 1960, effective 2001. Appendix B2 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" affected all cohorts equally. The average SPI-relevant gross wage in 2000 was € 22,313 (or DM54,256 per year), see Social Code Book VI,
https://www.gesetze-im-internet.de/sgb_6/anlage_1.html?

## B1 Public DI Health Assessment and Eligibility

Details of the application procedure are in German Social Law and Deutsche Rentenversicherung (2018). Applicants apply at an SPI field office and submit medical records. An independent third-party physician reviews the case.[11] Forty-four percent of all public DI applications have been rejected; this share has remained stable since 2000 (Deutsche Rentenversicherung, 2023b).

Before 2001, public ODI existed. It provided "social status protection" and insured the risk of becoming incapacitated to work in the previous occupation or a similar one in terms of income, education, and skills, see Section 2.2. Specifically, occupational disability applies when health conditions reduce an individual's work capacity" 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)).

[11] Sometimes these are state-employed physicians (*Amtsärzte*), and sometimes they are regular specialists practicing in the county of residence of the applicant. They must not have a pre-existing relationship with the applicant.

Today, only public WDI exists. The main WDI eligibility criterion is whether applicants' health limitations prevent them from working three hours per day in *any* job (Appendix Figure B1, column four). If applicants' work capacity is less than three hours daily, full WDI is granted; if it is between three and six hours daily, partial WDI is granted (Deutsche Rentenversicherung, 2020). The reform changed how work capacity was medically assessed: from an earnings capacity test[12] to an hours capacity test as discussed. Benefits are re-certified every three years and become permanent after nine years. If work capacity is not expected to improve, permanent DI benefits can be granted earlier, which applies to half of all new cases.

[12] Pre-reform, the applicant must be unable to earn more than 640 DM (about €320 in 2001 or $480 today).

**Figure B2:** Standard Health Assessment Questionnaire of Private ODI Insurer

| | | | | |
|---|---|---|---|---|
| 42f | Nehmen oder nahmen Sie innerhalb der **letzten 10 Jahre** Drogen, verschreibungspflichtige Medikamente, Betäubungs-, Suchtmittel oder werden oder wurden Sie innerhalb der **letzten 10 Jahre** wegen der Folgen des Konsums von Alkohol beraten oder behandelt? | ☐ ja | ☐ nein | BAS |
| **Ärztliche Behandlung, Operationen und Krankhausaufenthalte** | | | | |
| 51a | Sind oder waren Sie in ambulanter Behandlung von Ärzten, Psychologen, Psychotherapeuten oder Angehörigen sonstiger Gesundheitsberufe (z.B. Krankengymnast, Heilpraktiker, Physiotherapeuten)? | ☐ ja | ☐ nein | BAS |
| 51b | Sind Sie **derzeit** oder waren Sie innerhalb der **letzten 5 Jahre** länger als 2 Wochen in Behandlung von Ärzten, Psychologen, Psychotherapeuten oder Angehörigen sonstiger Gesundheitsberufe (z.B. Krankengymnast, Heilpraktiker, Physiotherapeut)? | ☐ ja | ☐ nein | CA |
| 51c | Sind Sie in den **letzten 5 Jahren** durch **Ärzte** oder **andere Behandler** (z. B. Heilpraktiker, Psychotherapeuten) untersucht, beraten oder behandelt worden? | ☐ ja | ☐ nein | NAV, DIA, ALL |
| 51d | Sind oder waren Sie in den **letzten 5 Jahren** in psychotherapeutischer Behandlung | ☐ ja | ☐ nein | STG |
| 51e | Waren Sie in den **letzten 5 Jahren** wegen Beschwerden oder Krankheiten der Psyche, des Rückens, des Bewegungsapparats, des Herzens, des Kreislaufs oder einer Krebserkrankung in ärztlicher, physiotherapeutischer oder psychotherapeutischer Behandlung? | ☐ ja | ☐ nein | BAS |
| 52 | Wurden Sie in den **letzten 10 Jahren** wegen einer Sucht- bzw. Abhängigkeitserkrankung ärztlich beraten oder behandelt? | ☐ ja | ☐ nein | BAS |
| 53a | Erfolgten in den **letzten 5 Jahren** Operationen, Krankenhaus- bzw. Kuraufenthalte oder haben Sie einen Unfall, Verletzungen oder Vergiftungen erlitten? | ☐ ja | ☐ nein | CA |
| 53b | Haben in den **letzten 5 Jahren** Krankenhaus-, Rehabilitations-, Kuraufenthalte oder ambulante Operationen stattgefunden oder sind solche beabsichtigt oder ärztlich empfohlen? | ☐ ja | ☐ nein | CA, GEN |
| 53c | Haben in den **letzten 5 Jahren** Krankenhaus-, Rehabilitations-/Kuraufenthalte oder ambulante Operationen (z.B. Laserung der Augen, Athroskopie) stattgefunden oder sind solche für die nächsten 2 Jahre ärztlich empfohlen oder beabsichtigt? | ☐ ja | ☐ nein | STG |
| 53d | Wurden Sie in den **letzten 10 Jahren** in einem Krankenhaus-, Rehabilitations-, Kureinrichtungen untersucht, beraten, behandelt oder sind solche für die nächsten 12 Monate empfohlen oder beabsichtigt ? | ☐ ja | ☐ nein | DIA |
| 53e | Wurden Sie in den **letzten 10 Jahren** ambulante Operationen durchgeführt, z.B. an inneren Organen, Haut, Bändern oder Augen? | ☐ ja | ☐ nein | DIA |
| 53f | Haben in den **letzten 10 Jahren** Krankenhaus-, Rehabilitations-, Kuraufenthalte oder ambulante Operationen stattgefunden oder sind solche derzeit ärztlich empfohlen oder beabsichtigt? | ☐ ja | ☐ nein | BAS, ALL |
| **Sonstiges** | | | | |
| 61 | Haben Sie in den letzten 12 Monaten Zigaretten, Zigarren oder Pfeife geraucht, Schnupftabak oder Kautabak oder sonst Nikotin aktiv zu sich genommen? | ☐ ja | ☐ nein | CA |
| 62 | Haben Sie in den letzten 12 Monaten Zigaretten geraucht? Wenn ja, wie viele Zigaretten rauchen Sie täglich? | ☐ ja | ☐ nein | BAS, GEN |

*Source:* The figure shows a standard health assessment questionnaire by a German private ODI insurer. In addition to age, occupation, medical diagnoses and diseases, and smoking status, all outpatient healthcare visits of the past 5 years and all inpatient healthcare of the past 10 years.

## B2 Public DI Benefit Calculation

We illustrate how the 2001 reform affected public benefits via a simple simulation assuming a stylized employment history.[13] As explained in Section 2, public DI is a part of SPI. Therefore, we first describe the primary method of calculating statutory retirement benefits. Then, we explain how disability benefits are calculated.

German SPI uses a point system, where the gainfully employed earn pension points ($pp_{it}$) during their work lives. A pension point equals the ratio of *individual* gross labor income ($I_{it}$) to *average* gross 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 €). This point 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 and full WDI pensions. Since 2001, it has been 0.5 for partial WDI and ODI benefits. Moreover, a fourth factor accounts for actuarial deductions ($AD_i$) if people retire before the statutory retirement age, which was 65 at the time of the reform and is now 67. These 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 before the statutory retirement age, pensions solely based on prior contributions would be relatively low. Hence, the DI benefit calculation assumes a "reference age." For the period between the entry of work disability and this reference age, individuals' prior *average* pension points are applied, and the system assumes the individual would have continued working until the reference age.

Before 2001, the reference age was 55, and the years until age 60 were valued with $1/3 \times$ average pension points. For example, an employee who hit work disability at age 40 would

[13]The fundamental 2001 reform introduced a new Social Code Book IX, entitled "Rehabilitation and Participation of Disabled Persons" (*Rehabilitation und Teilhabe von Behinderter Menschen*). Before 2001, most of these regulations were part of *Schwerbehindertengesetz*.

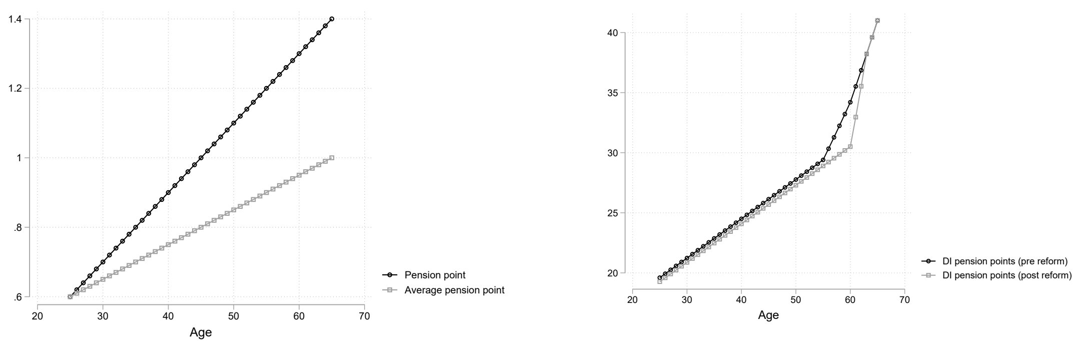
receive an additional 15 + 5/3 years of her average pension points. Before 2001, there were no actuarial deductions for WDI or ODI ($AD_i = 1$). $PT_i$ was 0.66 for ODI and 1 for WDI. Starting in 2001, $PT_i$ has been 0.5 for partial WDI as well as grandfathered ODI, and remained 1 for full WDI benefits.[14]

The 2001 reform increased the reference age to 60, but introduced actuarial deductions for years before 60. These deductions were capped at three years or 10.8% ($AD_i = 0.892$). As the large majority of disability shocks occur before age 57, the share of DI recipients with maximum deductions of 10.8% exceeded 90%.[15]

**Simulation.** Next, we simulate the effects of the 2001 reform on benefits for a stylized individual. We assume an increasing relative wage position approximately equal to one 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 B3a shows the average pension points by age.

**Figure B3:** 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 to (i) the grandfathered cohorts who could still claim public ODI, and (ii) the newly introduced partial WDI benefits for people who can work more than three but less than six hours a day in any job.

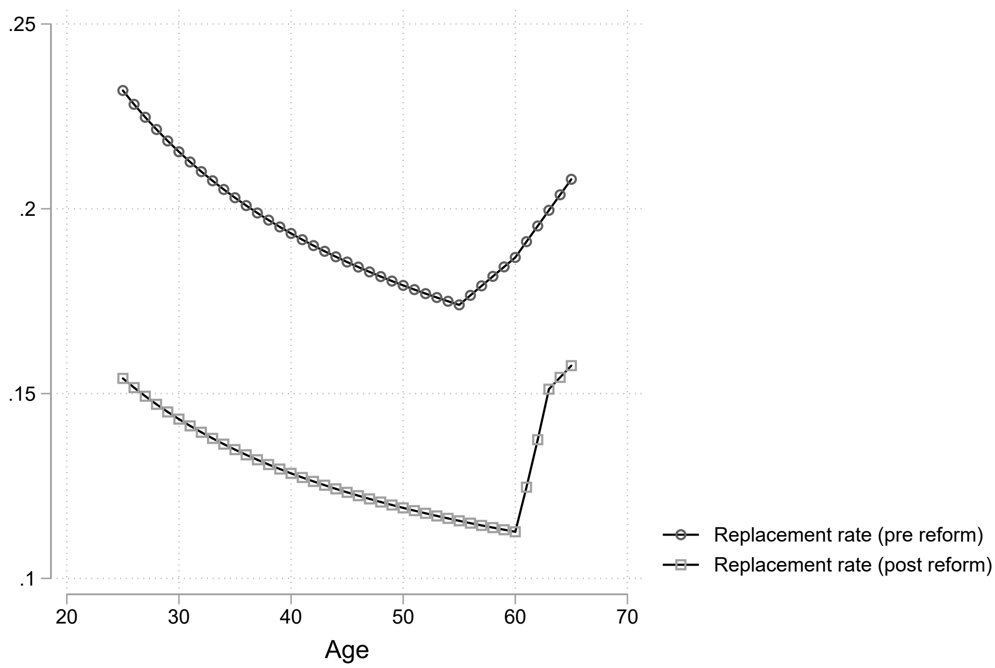
The introduction of actuarial deductions and the increase of the reference age to 60 approximately cancel each other out for most ages. Figure B3b shows that total pension points are slightly lower in the post-reform period. The largest difference applies between ages 56 and 61.

[14]If reasonable part-time work was unavailable, full WDI benefits could be granted (Viebrok, 2018).

[15] Several studies document a high poverty risk among WDI recipients (Krause et al., 2013; Geyer, 2021; Becker et al., 2023). Consequently, policymakers increased WDI benefits again by increasing the "reference age" for new recipients to 62 in July 2014, and to 65 years and eight months in 2019. It now equals the statutory retirement age and will increase to 67 years by 2031. Similarly, the age threshold for actuarial deductions has been raised.

Next, we calculate replacement rates by age for a single without additional income. To do so, we divide disability benefits by labor income. Figure B4 shows public ODI replacement rates in the pre- and post-reform periods. Before 2001, replacement rates were highest at 0.23 at age 25, then decreased linearly to 0.17 until the reference age of 55, after which they increased sharply again. After 2001, this pattern remained unchanged, but benefits are shifted downwards with lower replacement rates throughout, ranging from 0.11 to 0.16. Note that these benefit reductions solely applied *for the grandfathered cohorts* and for partial WDI, that is, employees with work capacity between three and six hours a day in any job. At the mean age of 46, when work disability typically occurs, the stylized replacement rate was 0.18 (pre-reform) and 0.12 (post-reform).

**Figure B4:** Replacement Rates Pre and Post-Reform




*Source:* own illustration. Note that the post-reform benefits apply to (i) the grandfathered cohorts who could still claim public ODI, and (ii) the newly introduced partial WDI benefits for people who can work more than three but less than six hours a day in any job.

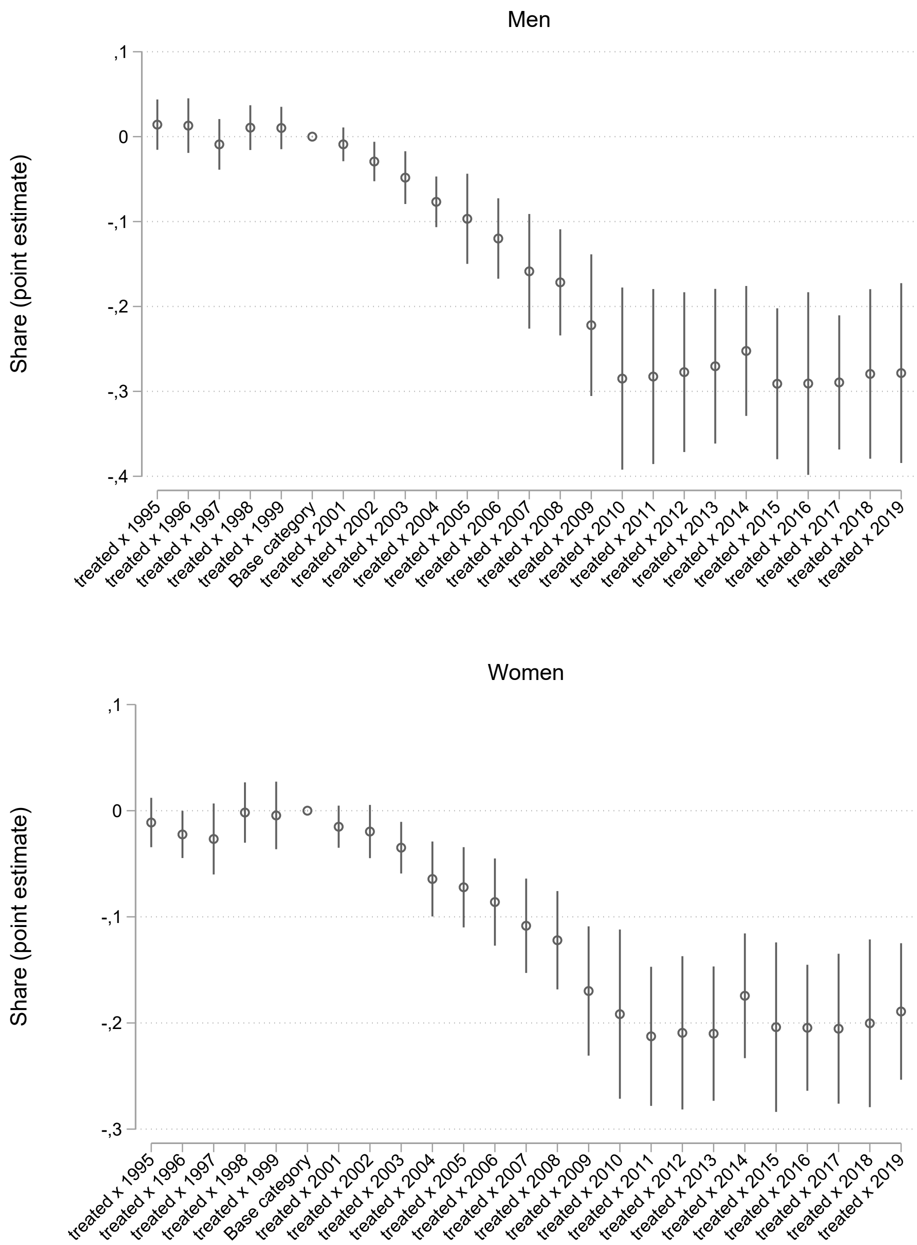
As a robustness check and validation, we use individuals' expected, self-reported pension replacement rate from the representative SAVE. As we lack precise individual work histories, summarizing individual-level information succinctly allows us to exploit this knowledge and expectations about future pensions. 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. 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 (b) current net household income (average 12%). We use the latter as the default when we simulate the reform effect in our model, but the results are similar when using the former.

Reassuringly, both the simulation above and the calculation based on self-reported replacement rates produce very similar results.

# C Impact of 2001 Reform on Public DI

## C1 Impact of 2001 Reform on Public DI Inflows (Administrative Data)

**Figure C1:** 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.

**Table C1:** Impact of 2001 Reform on Public DI Inflows (Administrative Data, by Gender)

| **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 controls 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 the main text for more details.

## C2 Impact of 2001 Reform on Public DI Case Load (Survey Data)

This section validates our first-stage findings in the main text using representative household data from the German Socio-Economic Panel Study (SOEP) and an alternative identification approach. The SOEP enables us to observe representative samples of each cohort, not just inflows, unlike the administrative data. Goebel et al. (2019) provides more details on the SOEP.

**Sample Selection.** We select years 1995 to 2016 and respondents between the ages 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, but routinely omit birth cohorts 1960 and 1961 to exclude potential transition and anticipation effects. Table C2 shows summary statistics, with our primary outcome variables in the upper panel and the covariates in the lower panel.[16]

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

$$y_{it} = \alpha + \beta D_i + \psi(1 - D_i) f(z_i - c) + \gamma D_i f(z_i - c) T_t + X'_{it}\tau + \delta_t + \rho_s + \epsilon_{it} \tag{12}$$

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 ($\delta_t$) and state ($\rho_s$) fixed effects. $X'_{it}$ represents a rich set of socio-demographic, educational, and job-related control variables as listed in Table C2. For example, the average age is 45, 52% are women, and 71% are married. Approximately 20% have completed 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 $\epsilon_{it}$ (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). The main results are that we present robust and bias-corrected estimates (Calonico et al., 2018). We also vary the bandwidth, use data-driven bandwidth selection (Calonico et al., 2020), and covariates (Calonico et al., 2019).[17] Moreover, our estimates are robust to implementing methods for discrete running variables following (Kolesár and Rothe, 2018).

The central RD identification assumption implies that no other factor would have affected public DI caseload trends discontinuously at the birth year level. We are unaware of another reform or factor that could invalidate this assumption; the appendix provides further evidence that other covariates trend smoothly at the cutoff $c$.

**Outcome.** The SOEP Group provides a time-consistent longitudinal binary variable that

[16] We provide summary statistics on the full sample, including birth years 1960 and 1961, as we use the full sample to produce empirical moments (Figure E4) as inputs for the model.

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

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*.[18] According to Table C2 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 B1 very well.

**Results.** Figure C2 plots public DI recipiency rates by birth cohorts for the pre-reform years in the left column (as a placebo test), and the post-reform years in the right column. Furthermore, the figure utilizes two distinct SOEP measures of DI receipt. The graphs show unconditional scatter by birth year, overlaid with polynomial quadratic smoothing plots. The visual evidence corroborates the findings from the administrative data: we see an apparent discontinuous decrease in the probability of receiving a public DI pension for the treated cohorts in post-reform years. Note that the DI level is higher for post-2001 years, as our respondents are older than those from 1995 to 2000. The decreasing slopes imply decreasing DI rates by birth cohort.

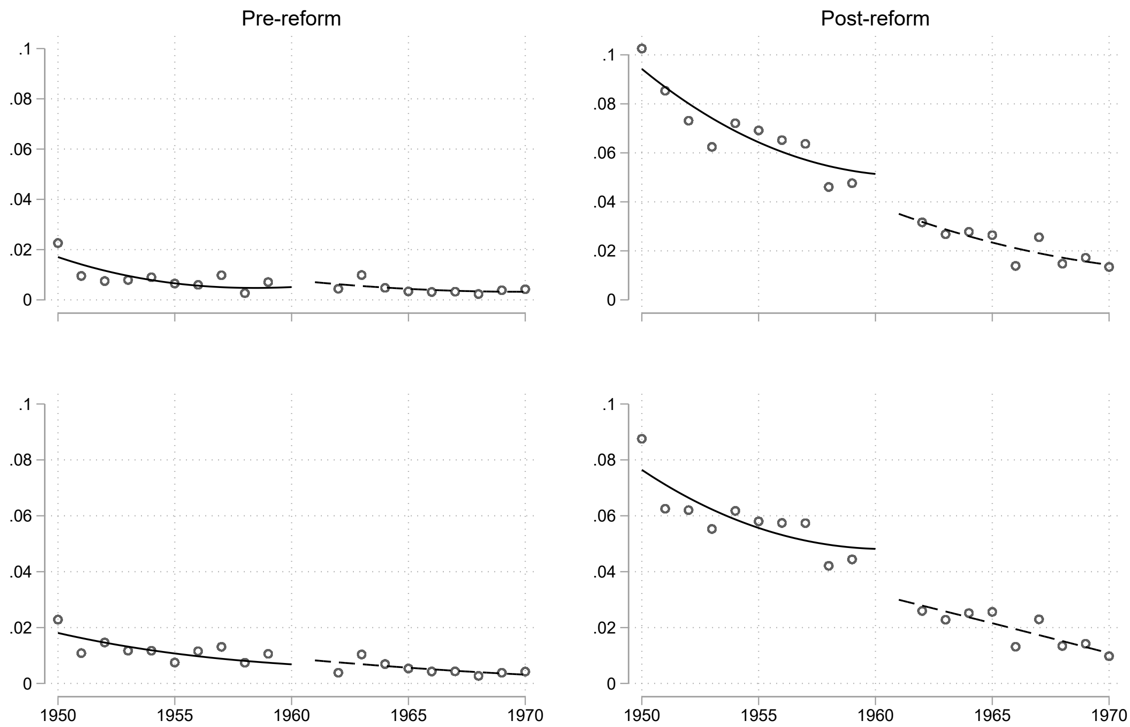
Table C3 shows the RD results using local polynomial RD methods for the post-reform period from 2001 to 2016. The column headers indicate the outcome measure; the lower panel adds socio-demographic and educational covariates as indicated. The models in columns (3) and (4) use *Public DI I* but restrict 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 shown, we obtain statistically significant results for 22 out of 24 models; all 24 models yield consistently negative point estimates, in line with Figure C2. 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 nontreated cohorts, 6.7%, the latter estimates translate into a decrease of 22%. The reduction in size for households with one member is very similar, whereas the decline for non-married individuals is even larger. Overall, the findings confirm and validate the results from administrative data that focus on inflows.

[18] 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.

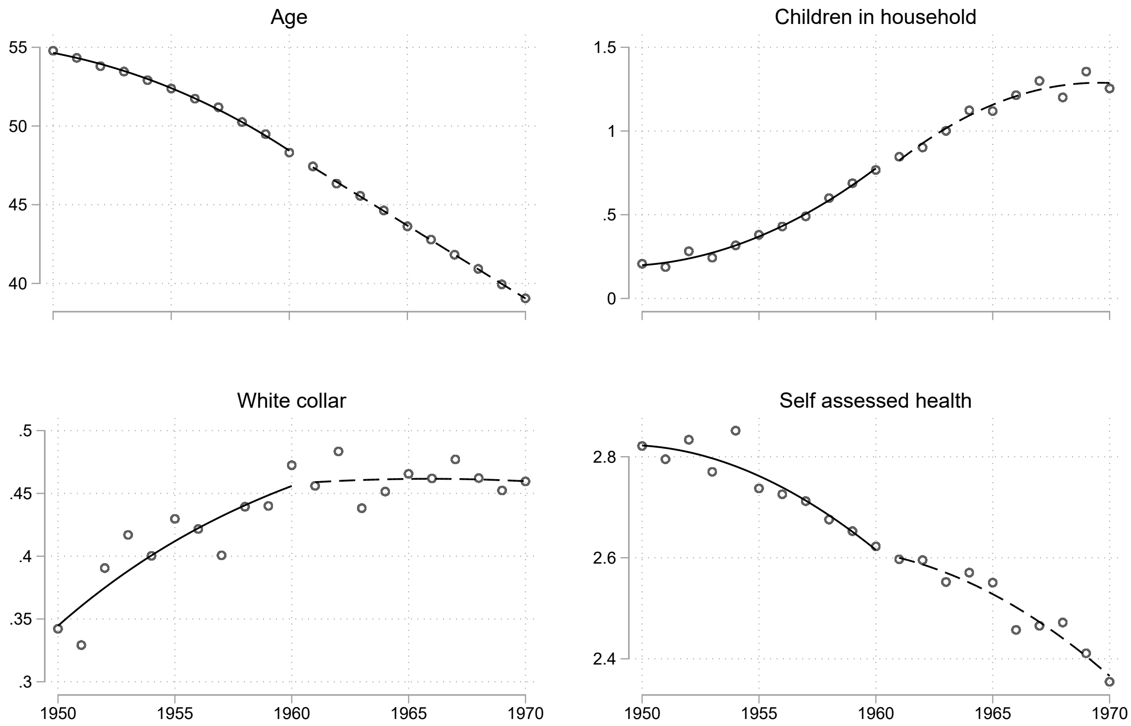
Robustness checks in Cao et al. (2022) vary the bandwidth and polynomials, the weights, run donut RD models, and add a set of covariates (Calonico et al., 2019, 2020; Pei et al., 2022); Figure C3 shows that covariates such as age, children in the household, white collar, or Self-Assessed Health (SAH) trend smoothly at the cutoff 1961. Figure C4 presents a density plot of the running variable, as described in citettech McCrary2008.

**Figure C2:** Effect of 2001 Reform on Public DI Using Representative SOEP Data



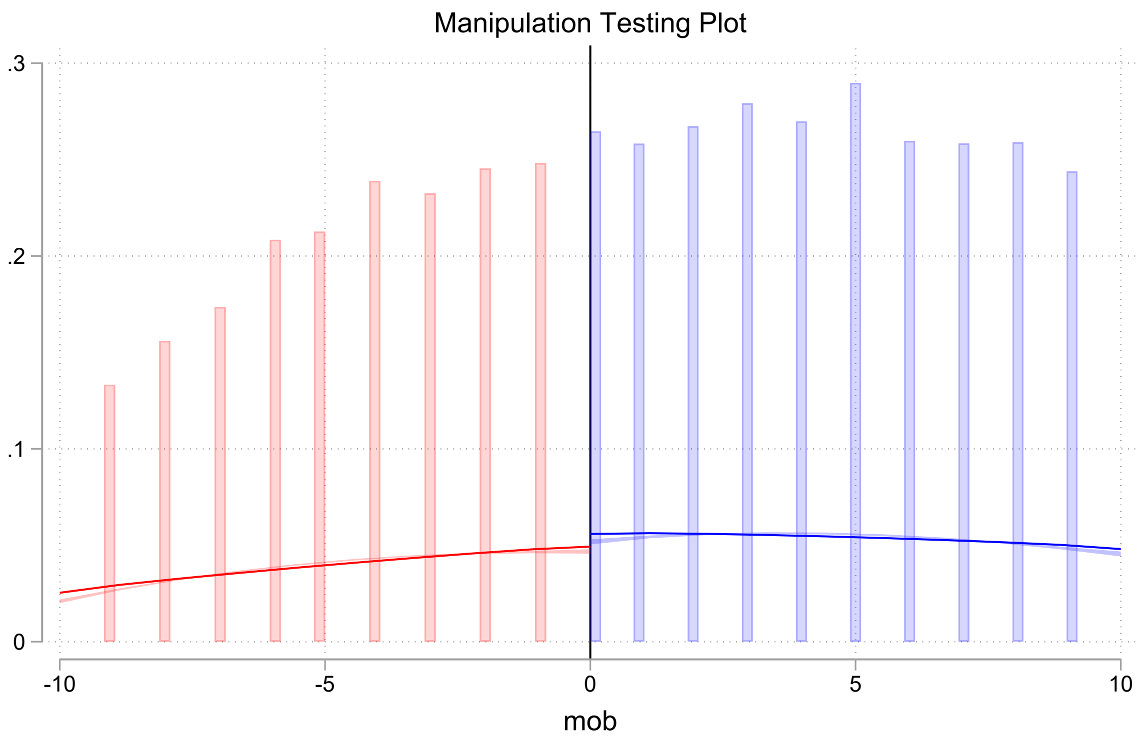

*Source:* SOEP v.33 – 95% sample. The left column shows the pre-reform years, and the right column shows the post-reform years. The first row shows *Public DI I*, and the second row shows *Public DI II*. All figures show the raw nonparametric means of public disability receipt by birth year, overlaid with separate quadratic trends before and after the cutoff. Other robustness checks vary the bandwidth and polynomials, the weights, run donut RD models and add covariates (Cao et al., 2022); study the smoothness of covariates (Figure C3), and carry out density plots of the running variable (Figure C4, McCrary (2008)).

**Figure C3:** 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 include varying the sample and indicator (Figure C2), varying the bandwidth and polynomials, the weights, running donut RD models, and adding covariates (Cao et al., 2022). Additionally, density plots of the running variable are provided (Figure C4).

**Figure C4:** Effect of 2001 Reform—Density Plot




*Source:* SOEP v.33 – 95% sample. The figure shows a density plot of the running variable for RD models similar to equation (12), estimated using local polynomial regressions with quadratic polynomials and univariate weights. Other robustness checks vary the sample and indicator (Figure C2), the bandwidth and polynomial (Cao et al., 2022), study the smoothness of covariates (Figure C3), and carry out density plots of running variables (Figure C4, McCrary (2008)).

**Table C2:** Descriptive Statistics, SOEP Data, 1995-2016

| | Mean | SD | Min | Max | N |
|---|---|---|---|---|---|
| **Panel A. Outcomes** | | | | | |
| Public DI I | 0.0329 | 0.1783 | 0 | 1 | 157783 |
| Public DI II | 0.0289 | 0.1675 | 0 | 1 | 157783 |
| Severe health limitation | 0.0554 | 0.2287 | 0 | 1 | 75638 |
| | | | | | |
| Non employed | 0.1870 | 0.3899 | 0 | 1 | 157783 |
| Full-time employed | 0.5949 | 0.4909 | 0 | 1 | 157783 |
| Individual total income (equivalized) | 28118.5884 | 29946.6603 | 0 | 2580000 | 157783 |
| Subjective Well-Being | 6.9186 | 1.7773 | 0 | 10 | 157783 |
| **Panel B. Socio-demographics** | | | | | |
| Age | 44.5266 | 7.7593 | 25 | 59 | 157783 |
| Female | 0.5230 | 0.4995 | 0 | 1 | 157783 |
| Married | 0.7088 | 0.4543 | 0 | 1 | 157783 |
| Single | 0.1296 | 0.3359 | 0 | 1 | 157783 |
| Children in household | 0.9140 | 1.0669 | 0 | 10 | 157783 |
| Adults in household | 0.3603 | 0.6713 | 0 | 7 | 157783 |
| Household size | 1.2743 | 1.1669 | 0 | 12 | 157783 |
| Dropout | 0.0234 | 0.1511 | 0 | 1 | 157783 |
| Schooling 9 yrs | 0.2573 | 0.4371 | 0 | 1 | 157783 |
| Schooling 10 yrs | 0.3620 | 0.4806 | 0 | 1 | 157783 |
| Schooling 13 yrs | 0.2012 | 0.4009 | 0 | 1 | 157783 |
| Civil servant | 0.0583 | 0.2343 | 0 | 1 | 157783 |
| Self-employed | 0.0946 | 0.2927 | 0 | 1 | 157783 |
| White collar | 0.4219 | 0.4939 | 0 | 1 | 157783 |
| Public sector | 0.2087 | 0.4063 | 0 | 1 | 157783 |
| Part-time employed | 0.2144 | 0.4104 | 0 | 1 | 157783 |
| In job training | 0.0025 | 0.0497 | 0 | 1 | 157783 |

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

**Table C3:** Impact of 2001 Reform on Public DI (Representative SOEP Data)

| **Panel A** | *Public DI I* | *Public DI II* | Non-Married | Single Households |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| 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) |
| Conventional | -0.012*** | -0.014*** | -0.005 | -0.016** |
| | (0.0038) | (0.0037) | (0.0086) | (0.0077) |
| Year FE | yes | yes | yes | yes |
| State FE | yes | yes | yes | yes |
| Age & Gender | 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 under 60 and those from birth cohorts between 1950 and 1970 are included. See Goebel et al. (2019) for more details about the SOEP. The tables display the point estimates obtained using local polynomial regressions, similar to equation (12) (Calonico et al., 2014, 2017, 2018, 2019, 2020), with a bandwidth of ten, a univariate kernel, and a quadratic polynomial. Column (2) shows results for an alternative *Public DI II* measure. Column (3) selects non-married respondents, and column (4) selects single households. Other robustness checks show results for the pre-reform period (Figure C2), vary the bandwidth and polynomials (Cao et al., 2022), study the smoothness of covariates (Figure C3), and carry out density plots of running variables (Figure C4).

## C3 Pre- and Post-Reform Consequences of a Health Shock.

In this subsection, we shed light on the question: For the treated cohorts without access to public ODI, how does a health shock materialize compared to the non-treated cohorts, given other social insurance strands and intra-household risk sharing? Table C4 uses SOEP panel data from 2001 to 2016 and runs simple OLS models with year and state fixed effects, and socio-demographic controls. 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), and 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) pre-tax income, and (4) her subjective well-being.

As seen in Table C4, the onset of a severe health limitation almost doubles the likelihood of taking up public DI in the next year (column (1)) and, by the same share of 9 ppt, increases non-employment. Further, total annual income decreases significantly by €4.1K (-12%) 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 Figure C2 and Table C3 (column (1)), it is imprecisely estimated. Similarly, the interaction effects suggest (imprecise) increases in non-employment of about four percentage points (column (2)) and small, insignificant effects for changes in income (column (3)) and well-being (column (4)). Overall, the findings for total individual income (including social insurance benefits) in column (3) show how the German social insurance system absorbs the financial consequences of a work-limiting health shock.

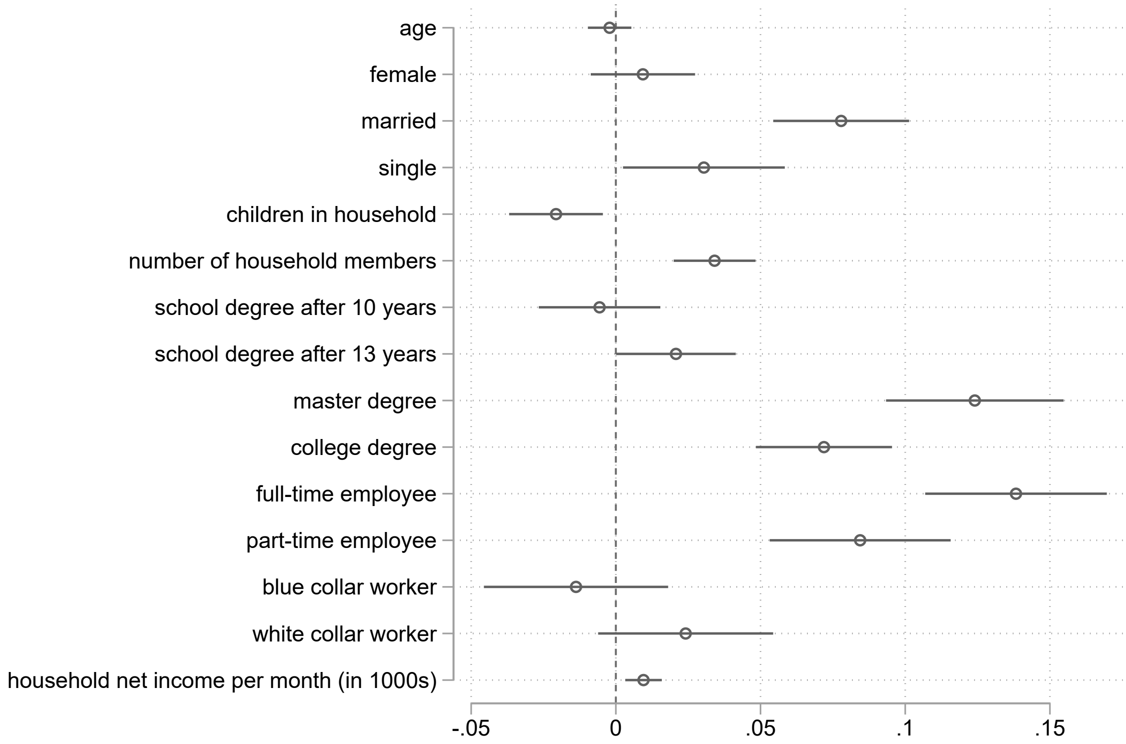
**Table C4:** Consequences of a Health Shocks: Treated vs. Nontreated

| | 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 | -0.0056 | -0.2161 | -17,365 | -1.6655** |
| | (0.0274) | (0.3367) | (14,193) | (0.6866) |
| Mean | 0.0404 | 0.1780 | 35,599 | 7.08 |
| N | 45,571 | 45,571 | 45,571 | 45,446 |
| $R^2$ | 0.0593 | 0.0314 | 0.0469 | 0.0094 |
| Year + State FE | yes | yes | yes | yes |
| Socio-demographic | yes | yes | yes | yes |
| Education | yes | yes | yes | yes |

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

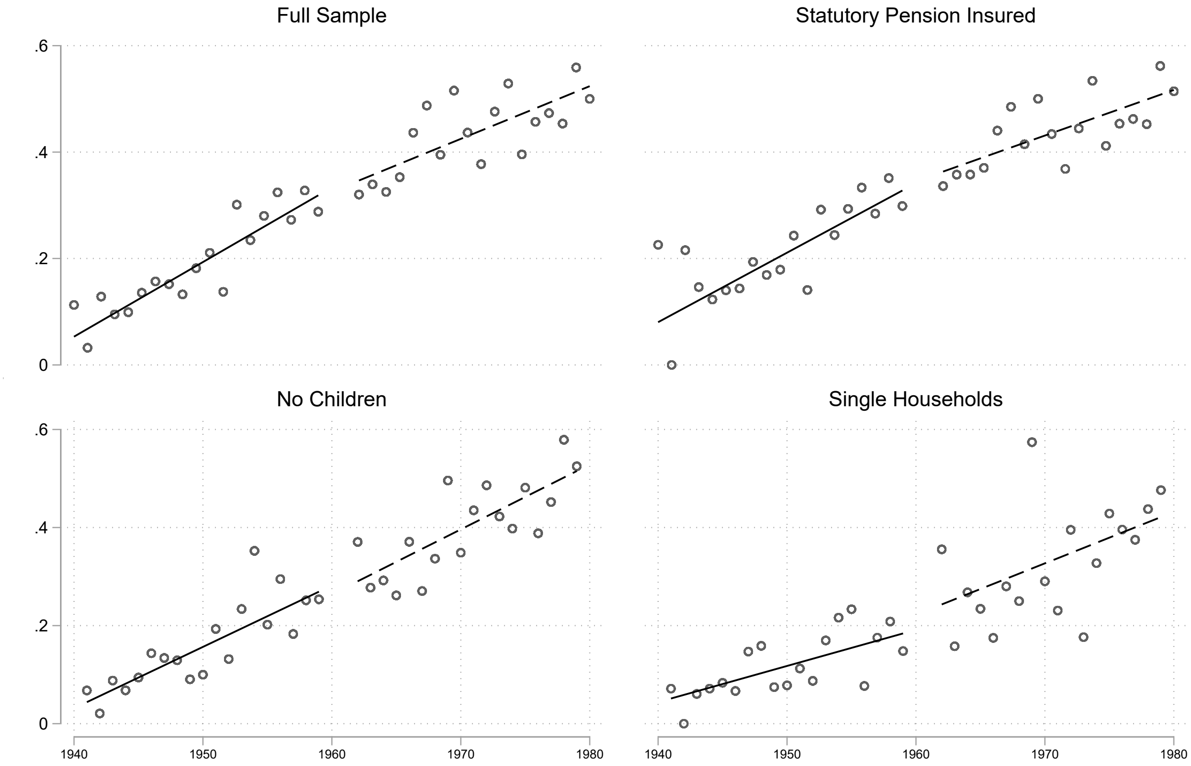
# D Impact of 2001 Reform on Private ODI Take-Up (Survey Data)

**Figure D1:** Socio-Demographic Predictors of Private ODI Coverage



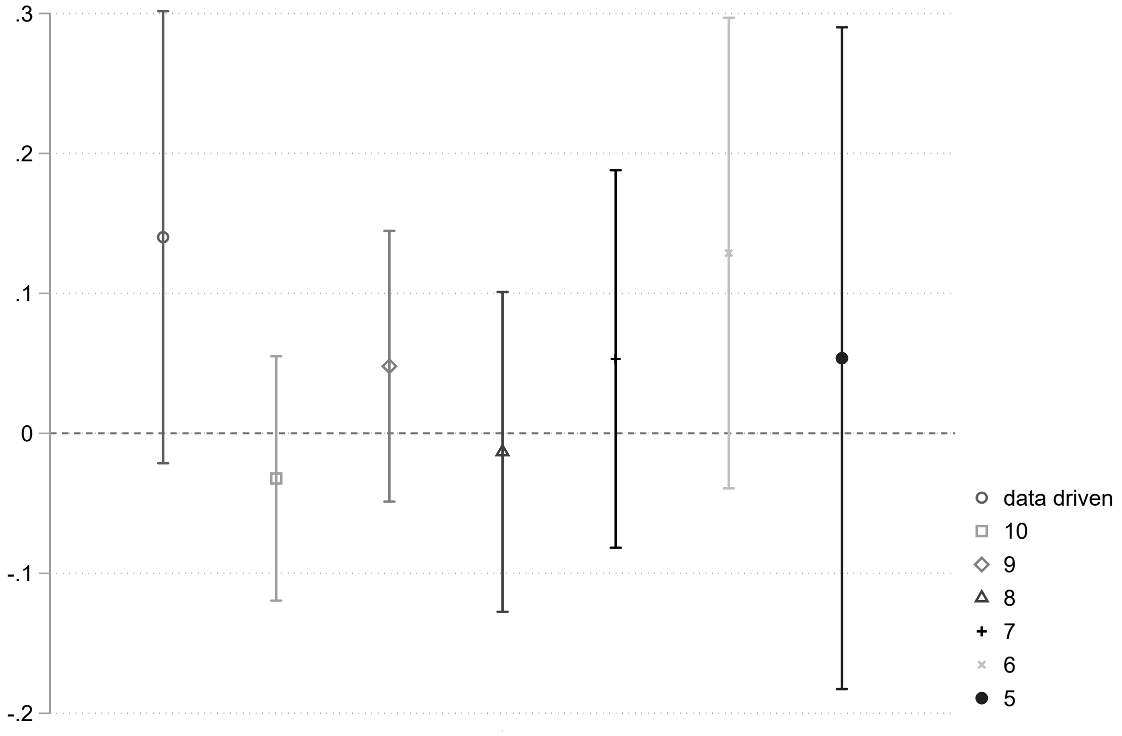

*Source:* The figure shows coverage predictors for private ODI policies. They stem from a multivariate regression and include the self-employed and civil servants (with civil servants omitted as the baseline category). Other omitted baseline categories include individuals who did not work at the time of the interview, those without a bachelor's or master's degree, and those with a degree obtained after 9 years of schooling, or those who dropped out.

**Figure D2:** Effect on Private ODI Coverage Using Representative SAVE Data



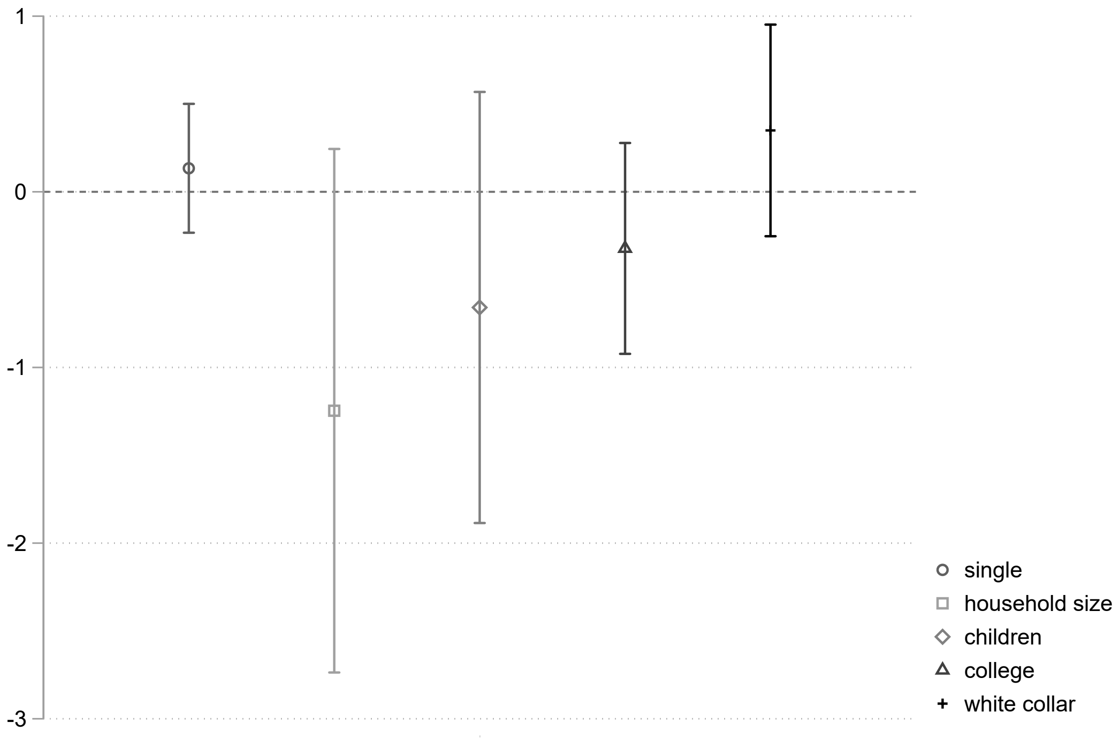

*Source:* SAVE data 2001-2010. The figures show the raw nonparametric means of private ODI coverage by birth year, overlaid with separate linear trends before and after the cutoff. The upper left graph is the default Figure 3; the upper right figure focuses on those eligible for Public DI; the bottom left focuses on the childless; and the bottom right focuses on one-person households. Other robustness checks vary the bandwidth (Figure D3), study discontinuities in covariates (Figure D4), carry out density plots of the running variable (Figure D5), and vary polynomials as well as run donut RDs (Figure D6).

**Figure D3:** Effect on Private ODI Coverage—Local Polynomial RD Varying Bandwidth



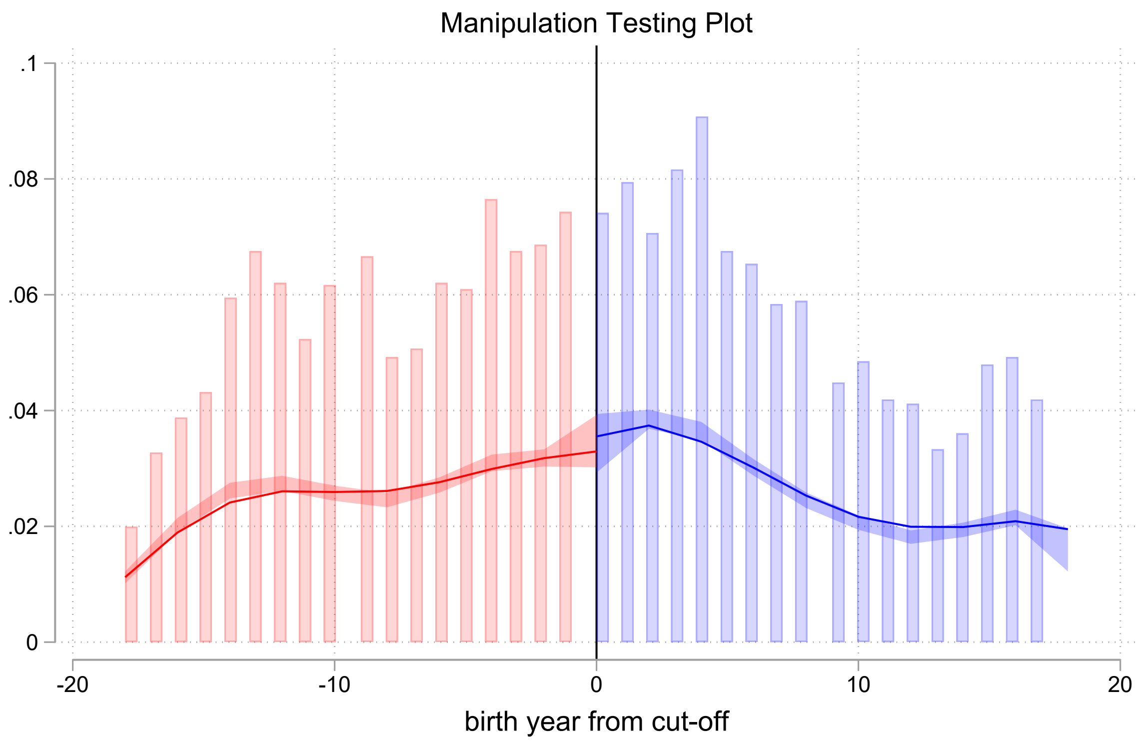

*Source:* SAVE data 2001-2010. The figures show point estimates of robustness checks varying the bandwidths of RD models similar to equation (12), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al., 2014, 2017, 2018). Other robustness checks vary the sample (Figure D2), study discontinuities in covariates (Figure D4), carry out density plots of running variables (Figure D5), and vary polynomials as well as run donut RDs (Figure D6).

**Figure D4:** Effect on Private ODI Coverage—Discontinuities in Covariates



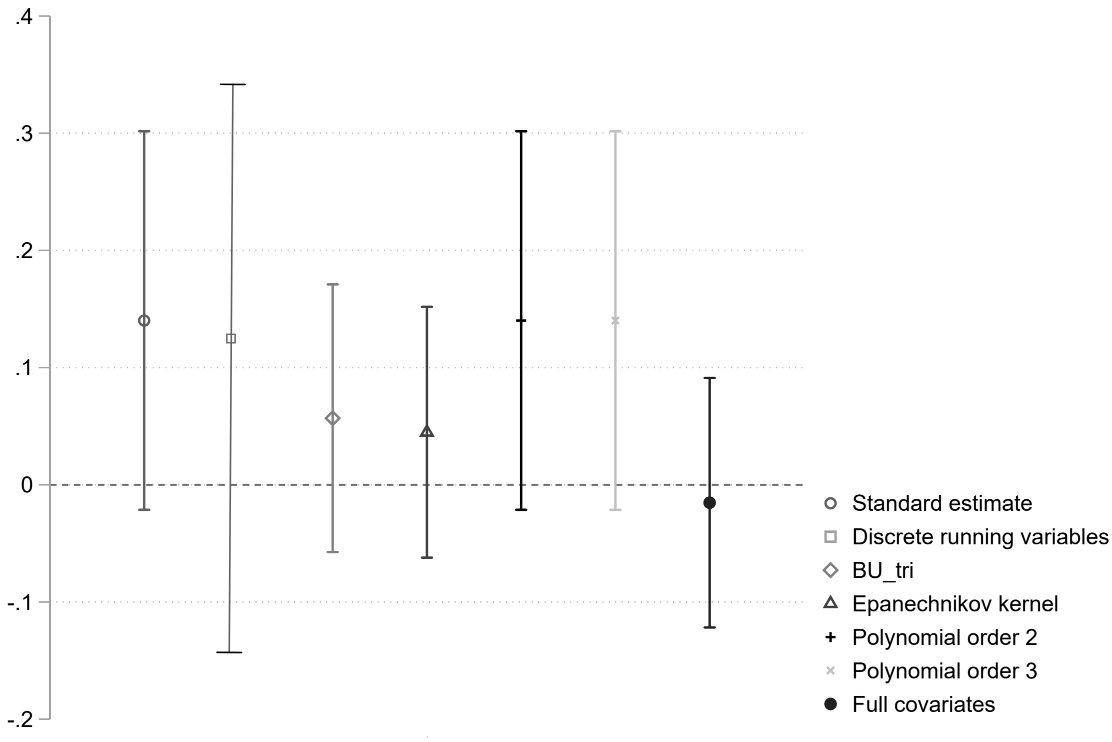

*Source:* SAVE data 2001-2010. The figures display point estimates from robustness checks that test for discontinuities in covariates using RD models similar to Equation (12), estimated via local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al., 2014, 2017, 2018). Other robustness checks vary the sample (Figure D2), vary the bandwidth (Figure D3), carry out density plots of running variables (Figure D5), and vary polynomials as well as carry out donut RDs (Figure D6).

**Figure D5:** Effect on Private ODI Coverage—Density Plot



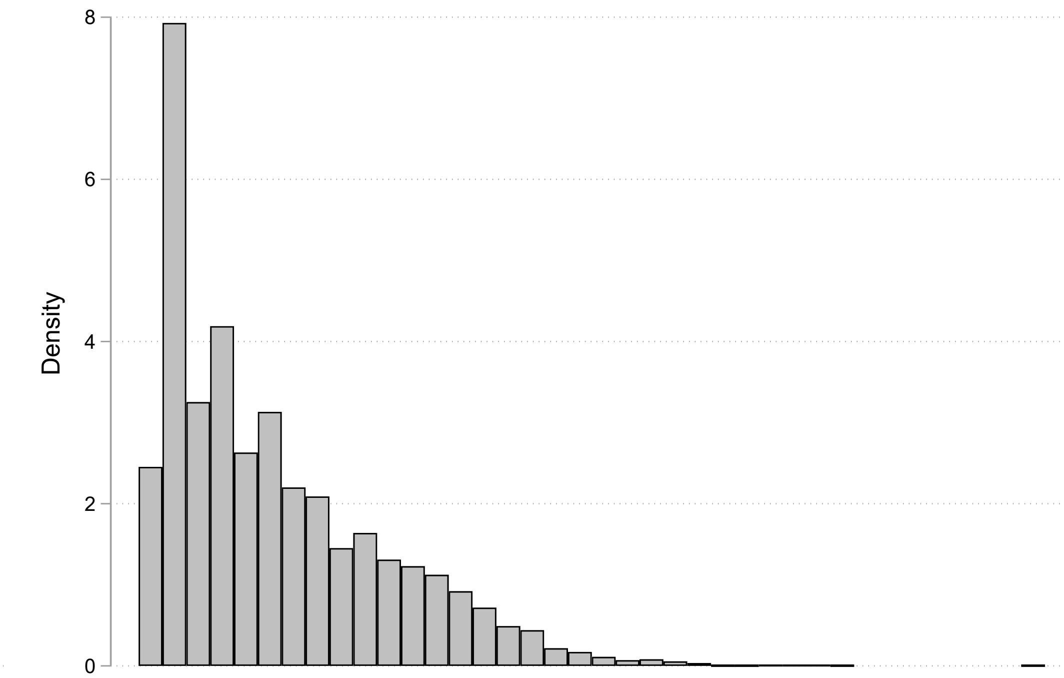

*Source:* SAVE data 2001-2010. The figure displays a density plot of the running variable for RD models similar to Equation (12), estimated using local polynomial regressions with quadratic polynomials and univariate weights (Calonico et al., 2014, 2017, 2018). Other robustness checks vary the sample (Figure D2), vary the bandwidth (Figure D3), carry out density plots of running variables (Figure D5), and vary polynomials as well as run donut RDs (Figure D6).

**Figure D6:** Effect of 2001 Reform: Local Polynomial RD—-Further Robustness




*Source:* SAVE data 2001-2010. The figure shows the point estimates of a robustness check, using methods for discrete running variables (Kolesár and Rothe, 2018), varying the order of the polynomials, varying weights, and adding covariates in RD models similar to equation (12), estimated via local polynomial regressions (Calonico et al., 2014, 2017, 2018, 2019). Other robustness checks vary the sample (Figure D2), vary the bandwidth (Figure D3, (Calonico et al., 2020)), study discontinuities in covariates (Figure D4), and carry out density plots of running variables (Figure D5, McCrary (2008)).

**Figure D7:** Distribution of Health Risk Score




*Source:* SAVE data 2001-2010. The health risk score is generated using principal component analysis and subjective and objective health measures from SAVE.

**Table D1:** Descriptive Statistics, SAVE Data, 2001-2010

| | Mean | SD | Min | Max | N |
|---|---|---|---|---|---|
| **Panel A. Key variables** | | | | | |
| Private ODI | 0.3239 | 0.4680 | 0 | 1 | 10721 |
| Expects Retirement Pre-60 | 0.0260 | 0.1592 | 0 | 1 | 10721 |
| **Panel B. Socio-demographics** | | | | | |
| Age | 45.35 | 10.26 | 22 | 62 | 10721 |
| Female | 0.5153 | 0.4998 | 0 | 1 | 10721 |
| Married | 0.669 | 0.4706 | 0 | 1 | 10721 |
| Single | 0.1794 | .3837 | 0 | 1 | 10721 |
| Children in household | 0.8714 | 1.0449 | 0 | 8 | 10721 |
| Household size | 2.656 | 1.2654 | 1 | 13 | 10721 |
| High school degree | 0.3353 | 0.4721 | 0 | 1 | 10721 |
| Master degree | 0.2684 | 0.4432 | 0 | 1 | 10721 |
| College degree | 0.3962 | 0.4891 | 0 | 1 | 10721 |
| Full-time | 0.4556 | 0.498 | 0 | 1 | 10721 |
| Part-time | 0.138 | 0.3449 | 0 | 1 | 10721 |
| Blue collar | 0.2101 | 0.4074 | 0 | 1 | 10721 |
| White collar | 0.4032 | 0.4906 | 0 | 1 | 10721 |
| Household net income (in 000s) | 2.429 | 2.442 | 0 | 120 | 10721 |
| **Panel C. Subjective and Objective Health** | | | | | |
| Health satisfaction 0-4/10 | 6.648 | 2.509 | 0 | 10 | 10721 |
| Concerns about own health | 0.2171 | 0.4123 | 0 | 1 | 10721 |
| SAH | 2.4553 | 0.8551 | 1 | 5 | 7811 |
| Serious Health Issues | 0.4672 | 0.4990 | 0 | 1 | 7811 |
| Heart disease diagnosed | 0.0648 | 0.2462 | 0 | 1 | 7811 |
| Stroke | 0.0193 | 0.1377 | 0 | 1 | 9580 |
| Chronic Lung Disease | 0.0604 | 0.2383 | 0 | 1 | 7811 |
| Cancer | 0.0454 | 0.2083 | 0 | 1 | 7811 |
| High Blood Pressure | 0.2284 | 0.4198 | 0 | 1 | 7811 |
| High Cholesterol | 0.1369 | 0.34378 | 0 | 1 | 7811 |
| # doctor visits | 0.6104 | 0.847 | 0 | 9 | 5706 |
| # days hospital | 0.2012 | 0.9268 | 0 | 27 | 5706 |
| Smoker | 0.359 | 0.4797 | 0 | 1 | 10721 |
| Normalized health risk score | 0.1547 | 0.125 | 0 | 1 | 5706 |
| **Panel D. Expectations and attitudes** | | | | | |
| Subj. life expectancy low | 0.1825 | 0.3863 | 0 | 1 | 9757 |
| Subj. life expectancy high | 0.1324 | 0.339 | 0 | 1 | 9757 |
| Savings 4 Unexp. Important | 0.6956 | 0.4602 | 0 | 1 | 9757 |
| Savings 4 Old Age Important | 0.7256 | 0.4462 | 0 | 1 | 9757 |
| No savings possible | 0.2108 | 0.4079 | 0 | 1 | 9757 |
| No savings, enjoy life | 0.0197 | 0.1389 | 0 | 1 | 9757 |
| Higher-income expected | 2.0909 | 2.9554 | 0 | 1 | 10721 |
| Inheritance expected | 0.0358 | 0.1858 | 0 | 1 | 10721 |

*Source:* SAVE data 2001-2010. Table shows the sample for the reduced-form RD estimate on take-up. Respondents below 60, civil servants and self-employed are omitted, as are birth cohorts 1960 and 1961; see main text for more details on sample selection. See Coppola and Lamla (2013) for more details about SAVE.

**Table D2:** Complier Analysis, SAVE Data, 2001-2005

| | Treated <46 | Treated <46 | Controls <51 | Controls <51 |
|---|---|---|---|---|
| **Panel A. Key variables** | | | | |
| Private ODI | 1 | 0 | 1 | 0 |
| | | | | |
| **Panel B. Socio-demographics** | | | | |
| Age | 34.34 | 34.10 | 46.71 | 46.83 |
| Female | 0.455 | 0.496 | 0.406 | 0.449 |
| Married | 0.663 | 0.543 | 0.812 | 0.697 |
| Single | 0.293 | 0.349 | 0.073 | 0.117 |
| Children in household | 1.124 | 1.028 | 1.209 | 1.024 |
| Household size | 2.925 | 2.697 | 2.940 | 2.668 |
| High school degree | 0.471 | 0.416 | 0.385 | 0.323 |
| Master degree | 0.231 | 0.148 | 0.265 | 0.183 |
| College degree | 0.705 | 0.687 | 0.650 | 0.687 |
| Full-time | 0.613 | 0.484 | 0.667 | 0.570 |
| Part-time | 0.132 | 0.144 | 0.158 | 0.123 |
| Blue collar | 0.213 | 0.225 | 0.226 | 0.226 |
| White collar | 0.465 | 0.344 | 0.479 | 0.393 |
| HH net income (000s) | 2.634 | 2.256 | 2.891 | 2.539 |
| | | | | |
| **Panel C. Subjective and Objective Health** | | | | |
| Health satisfaction 0-4/10 | 7.673 | 7.576 | 7.077 | 6.600 |
| Concerns about health | 0.072 | 0.109 | 0.111 | 0.191 |
| SAH | 2.078 | 2.090 | 2.208 | 2.570 |
| Serious health issues | 0.307 | 0.294 | 0.283 | 0.504 |
| Heart disease diagnosed | 0.009 | 0.025 | 0.038 | 0.058 |
| Stroke | 0.000 | 0.006 | 0.000 | 0.017 |
| Chronic lung disease | 0.005 | 0.023 | 0.000 | 0.033 |
| Cancer | 0.009 | 0.014 | 0.038 | 0.033 |
| High blood pressure | 0.046 | 0.076 | 0.113 | 0.182 |
| High cholesterol | 0.028 | 0.040 | 0.038 | 0.066 |
| Smoker | 0.367 | 0.453 | 0.389 | 0.464 |
| | | | | |
| **Panel D. Expectations and attitudes** | | | | |
| Subj. life expect. low | 0.123 | 0.140 | 0.150 | 0.177 |
| Subj. life expect. high | 0.188 | 0.134 | 0.192 | 0.146 |
| Savings 4 unexp. important | 0.725 | 0.649 | 0.705 | 0.642 |
| Savings 4 old age important | 0.808 | 0.644 | 0.812 | 0.683 |
| No savings possible | 0.107 | 0.276 | 0.090 | 0.238 |
| No savings, enjoy life | 0.013 | 0.018 | 0.009 | 0.016 |
| Higher income expected | 2.775 | 2.468 | 2.077 | 1.696 |
| Inheritance expected | 0.044 | 0.025 | 0.051 | 0.024 |

*Source:* SAVE data 2001-2005. Respondents below 60, civil servants, and self-employed are omitted. The table solely includes years up to 2005 to minimize deviations from the reform and limit the aging effect in our sample. See Coppola and Lamla (2013) for more details about SAVE.

# E Further Details on the Equilibrium Model

## E1 Optimal Contracts

This section summarizes optimal insurance contracts in the standard model with private information when adding administrative costs and allowing for a (means-tested) consumption floor. We rely 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 risks and at the bottom of the work incapacity risk distribution, $\theta^b$, or bad risks and at the top, $\theta^t$.

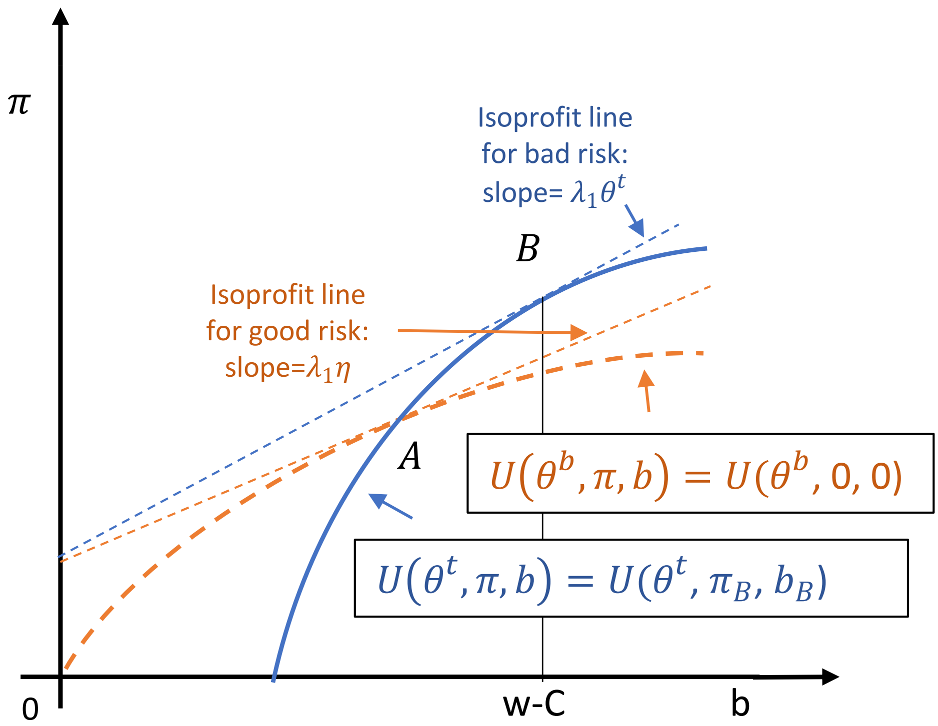
#### Standard Case: Just Private Information

The core of the standard case goes back to Rothschild and Stiglitz (1976) and Stiglitz (1977). Given the participation and incentive compatibility constraints, the insurer maximizes profits (see equation (6)). Figure E1 illustrates optimal contracts under the standard case. The x-axis shows the insured benefit $b$ and the costs of work incapacity, $w - C$, where $w$ represents the wage in the trained occupation and $C$ is the consumption floor. The y-axis shows the premium $\Pi$, increasing 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 can be seen, the bad risks have a higher marginal willingness to pay. The dashed curve intersecting at (0,0) represents the participation constraint when binding. 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 $(\Pi, b)$ that produce the same utility for individuals, given the participation and incentive compatibility constraints (both binding in the standard case).

Consequently, we obtain the optimal contract for the good risks where the flatter isoprofit

**Figure E1:** 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 incapacity distribution $\theta^b$ trading off premia ($\Pi$) 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 incapacity distribution $\theta^t$. 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.

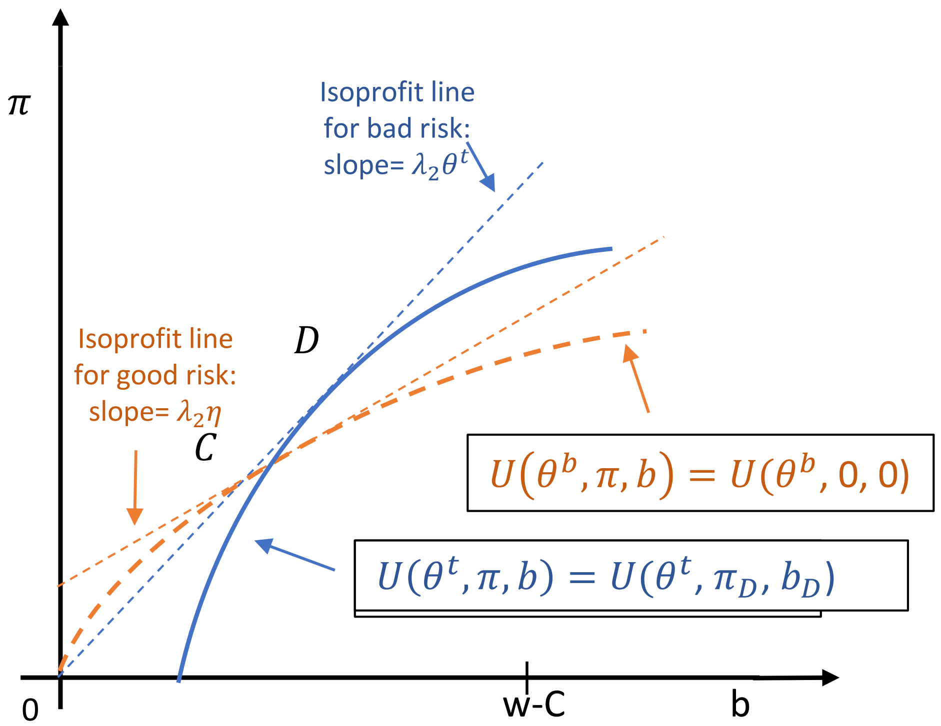
curve of the insurer touches the indifference curve of the good risks at point A. Compared to the optimal contract for the bad risks at B, the benefits and premium are lower; the contract solely provides partial insurance, whereas the optimal contract for the bad risks in B provides full insurance with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

As discussed, the standard case cannot produce coverage denials by insurers. Only the good risks can be voluntarily uninsured with (0,0) and make an ODI take-up that is not 100%. In other words, insurers always offer policies. Such a scenario can happen when the share of the population with low work incapacity risk, $\rho$, is small. Still, the dispersion of the true incapacity risk $\theta^i$—that is, unobserved by the insurer—is large. In this case, the good types are offered a profitable contract by the insurer but prefer to remain uninsured.

### 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) in practice, a rather unlikely case where only the bad risks are insured.

**Figure E2:** Optimal Contracts with Private Information and Admin Costs: Separating Equilibria



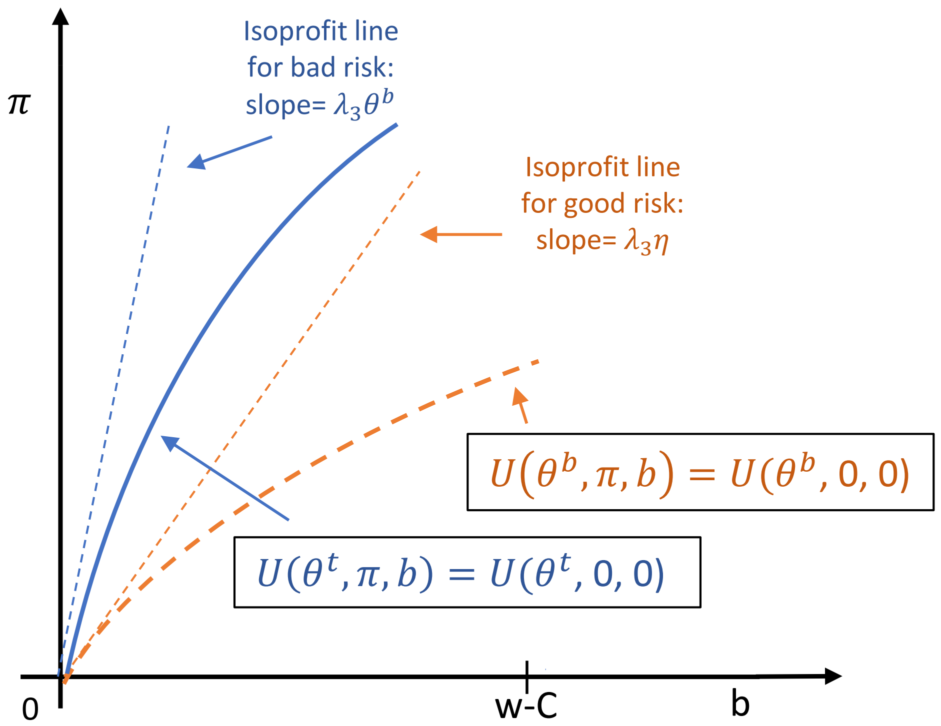

*Source:* The dashed indifference curve shows optimal contracts for good risks at the bottom of the work incapacity distribution $\theta^b$ trading off premia ($\Pi$) on the y-axis and coverage levels ($b$) on the x-axis. The solid indifference curve shows optimal contracts for bad risks at the top of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

Once variable administrative costs are introduced, optimal contracts for good and bad risks never provide full insurance. Furthermore, all members of a risk group may be denied coverage. These are the two relevant cases in practice. As shown in Figure E2, administrative costs result in steeper isoprofit curves for insurers. This implies that, in a separating equilibrium, the insurer offers policies with lower benefits and premiums. Hence, in Figure E2, optimal contracts for both groups provide less coverage and 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 E3 illustrates a

**Figure E3:** 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 incapacity distribution $\theta^b$ trading off premia ($\Pi$) 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 incapacity distribution $\theta^t$. 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.

scenario in which the entire risk group is denied coverage. This is because the insurer cannot offer a profitable contract with favorable coverage. The result is a pooling contract with (0, 0), and no one has insurance. Please see Chade and Schlee (2020) and Braun et al. (2019) for more details and formal proof.

#### Extended Case II: Private Information and Social Insurance

Braun et al. (2019) introduce an extension that includes a means-tested public insurer for long-term care costs ('Medicaid'). It crowds out private insurance benefits dollar by dollar. This is not the case in Germany, where private ODI benefits top up the basic WDI benefits. This implies that German private ODI also provides utility with public benefits, unlike in the U.S. case. Nevertheless, the main underlying mechanisms are the same in the German ODI case: public social insurance can lead to optimal contracts with partial coverage. Furthermore, it can result in the denial of coverage.

Public insurance generally increases individuals' utility in the case of no private insurance.

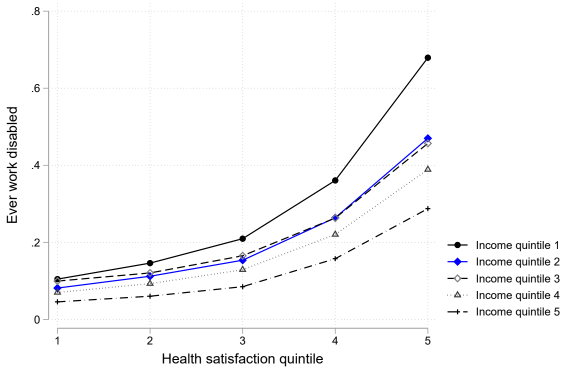
It thus reduces the demand for private insurance and the profits of private insurers. As it increases the individual's outside option, insurers lower premiums (to satisfy the participation constraint). However, if the consumption floor is high enough, the insurer cannot offer contracts that are still 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 is relevant in Germany, where the consumption floor is relatively high. As with administrative costs, whether an insurer denies coverage to entire risk groups depends on the dispersion of private information and the population share of the good risks $\rho$.

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 but more so for high-income populations whose income could drop more sharply. As explained, we use the representative SOEP to model the income shock distribution over the lifecycle and set the bounds for $\tau$ empirically (see Figure 6 and Table 2).

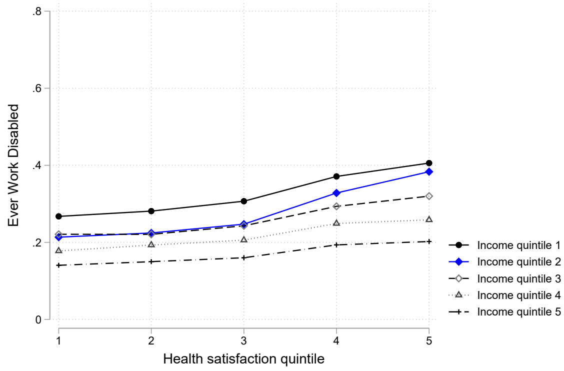
In conclusion, the equilibrium model includes multiple risk groups with observable $h, w, o$, whereas $\theta^i$ is private information. An ODI take-up rate of less than 100% is achieved through two channels. First, insurers deny coverage to entire groups. Second, some individuals are offered a profitable optimal policy, but those individuals prefer to self-insure.

## E2 Additional Figures and Tables

**Figure E4:** Lifecycle Risk of Work Incapacity by Income and Health Risk Score



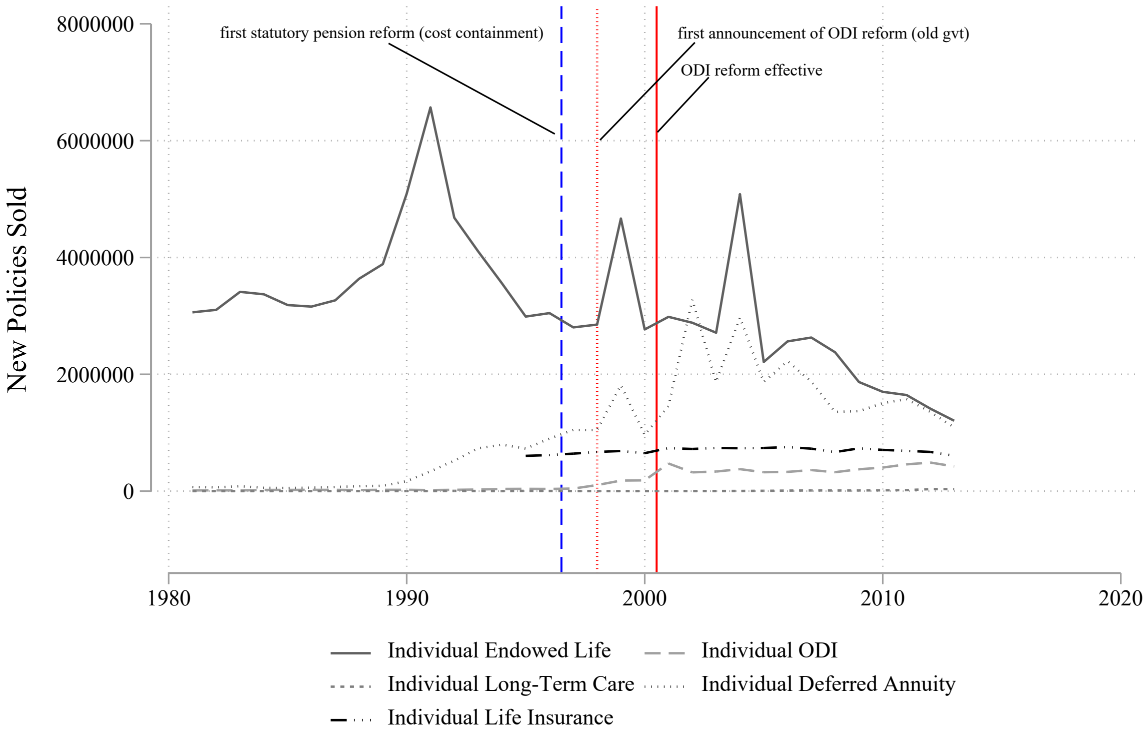

**(a)** unconditional




**(b)** conditional

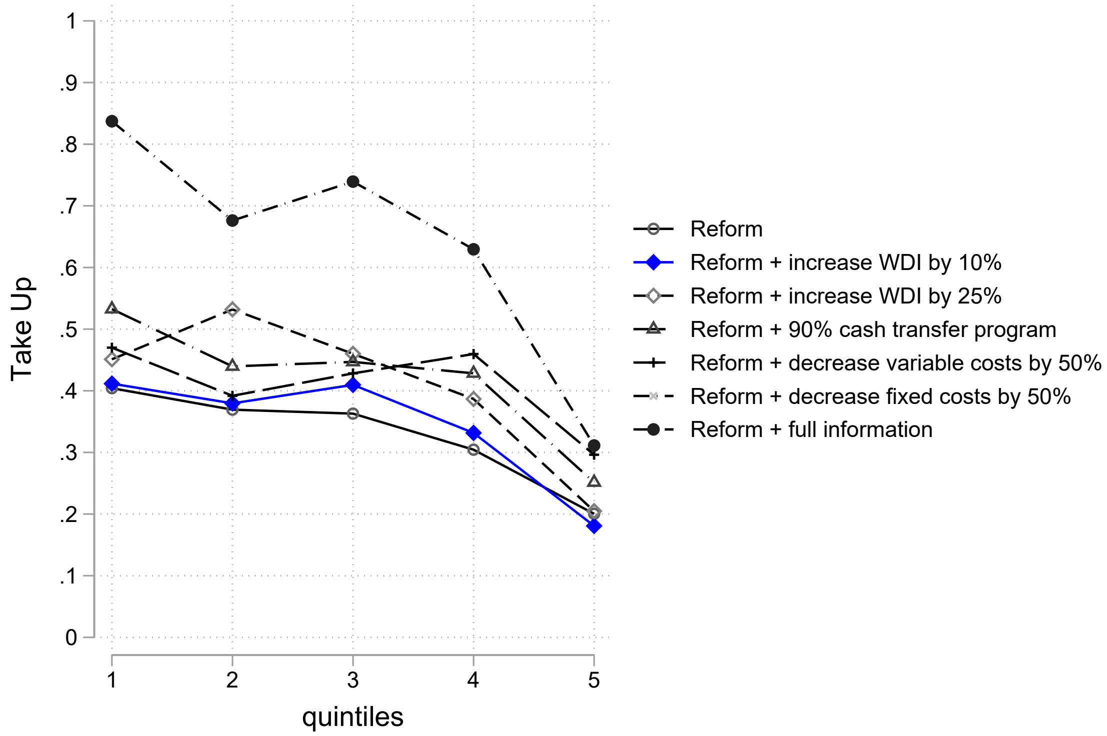
*Source:* SOEP v.33 – 95% sample. 1984-2016. Figure E4a plots the unconditional risk of a severe health limitation over the working ages by the health satisfaction quintiles and the five net household income quintiles. Figure E4b first regresses the lifecycle risk of severe health limitations on socio-demographics, 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. Unfortunately, the SOEP only includes health satisfaction (and the standard SAH measure) over the entire 33 SOEP years in our sample. The quintiles are not exact as they are derived from the 0-10 Likert scale.

**Figure E5:** Non-Group Insurance Policies Sold in Germany




*Source:* German Insurance Association (GDV), own illustration. The first vertical line (dashed blue) marks the passage of the Rentenanpassungsgesetz (1996), the first of a series of structural pension reforms. It tightened eligibility for early retirement and reduced benefits. The second vertical line (dotted red) indicates the time when the 2001 reform was first announced, still be the previous center-right government. The third vertical line (solid red) indicates when the 2001 reform took effect.

**Figure E6:** 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 (Figure D7). The other lines show take-up rates for (additional) policy simulations by health risk quintiles using the general equilibrium model (see Section 5).

**Table E1:** 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:* Table shows the average health risk score as in Figure D7 by income quintiles. The first row shows the empirical moments from SAVE, and the second row those produced by the model.

# F Parameter Calibration

We calibrate the key parameter $\psi$, the share of good types, and the lifecycle work incapacity probabilities by good and bad risk types $\{\theta^b_{h,w,o}, \theta^t_{h,w,o}\}$, using a grid search procedure that minimizes the distance between model-generated and empirical ODI take-up rates. The calibration employs a coarse-to-fine search strategy to efficiently identify the parameter value that best matches the empirical targets. The calibration strategy exploits a structural constraint: the weighted combination of type-specific probabilities must exactly reproduce the population-level work incapacity probabilities observed in the SOEP data:

$$\eta_{h,w,o} = (1-\psi)\cdot\theta^b_{h,w,o} + \psi\cdot\theta^g_{h,w,o} \tag{13}$$

where $\eta_{h,w,o}$ denotes the empirical population-level work incapacity probability for health quintile $h$, income quintile $w$, and age group $o$.

**Calibration Algorithm.** For each candidate value of $\psi$, the population constraint in the equation defines a mapping between the two sets of type-specific probabilities. We exploit this by using an iterative procedure that:

1. Fixes a candidate value of $\psi$
2. Initializes type-specific probabilities satisfying the population constraint ($\psi$ and the population constraint together span a possibility space for $\{\theta^b_{h,w,o}, \theta^t_{h,w,o}\}$)
3. Runs the full model to generate ODI take-up rates by health and income quintiles
4. Adjusts the type-specific probabilities using bisection search to reduce the distance between model-generated and empirical take-up rates

5. After each adjustment of one type's probabilities, recalculates the other type's probabilities to maintain the population constraint

6. Iterates 20 iterations per $\psi$ value

This nested algorithm ensures that at every iteration, the model perfectly matches the population-level work incapacity probabilities by construction, while searching for the combination of $\psi$ and type-specific probabilities that best match the empirical ODI take-up rates.

**Initial Grid Search.** We begin with the prior that $\psi \in (0.5, 1)$, reflecting the expectation that good types constitute a majority of the population, and conduct an initial coarse grid search over $\psi \in \{0.5, 0.6, 0.7, 0.8, 0.9\}$. Given the population-level work incapacity probabilities observed in the SOEP and an initial draw for $\psi$, we obtain the corresponding set of feasible combinations for the type-specific probabilities. We assess the model fit using the sum of squared deviations:

$$f(\psi) = ||\mathbf{r}_m(\psi) - \mathbf{r}_e||^2 \tag{14}$$

where $\mathbf{r}_m(\psi)$ denotes the vector of model-generated take-up rates by income and health status, and $\mathbf{r}_e$ the corresponding empirical targets. The coarse grid search indicates that $\psi = 0.7$ provides the closest match, suggesting that the optimal value lies within the interval $[0.6, 0.8]$.

**Fine Grid Search.** Building on these results, we conduct a refined search over a finer grid spanning $[0.6, 0.8]$, consisting of ten equally spaced points: $\psi \in \{0.6000, 0.6222, 0.6444, 0.6667, 0.6889, 0.7111, 0.7333, 0.7556, 0.7778, 0.8000\}$. For each value, we repeat the full iterative calibration procedure described above. The final calibrated parameter is $\psi = 0.6222$, which produces model-generated take-up rates with the smallest sum of squared deviations from the empirical targets while maintaining exact consistency with population-level work incapacity probabilities.

# G Consumption-Equivalent Variation

*Ex-ante* utility provides ordinal rankings of policy alternatives. We enhance this utility-based approach with consumption-equivalent variation (CEV), a widely used metric in quantitative economics. CEV expresses welfare changes as percentage changes in consumption. In our setting, CEV answers the question: *"By what percentage would we need to change individuals' income in period 1 in the counterfactual economy to make them equally well off as under the baseline economy?"* More formally, for an individual who earns income $w$ in the first period in the counterfactual economy, we find the percent increase in $w$ that satisfies:

$$EU^{Counterfactual}(w(1+x)) = EU^{Baseline}(w) \tag{15}$$

where $EU^{Counterfactual}(.)$ and $EU^{Baseline}(.)$ represent expected lifetime utility under each policy scenario, accounting for optimal savings and insurance decisions at each income level. $x$ stands for the individual-level CEV.

To provide an example: A CEV of 2.5% implies that individuals would need to be compensated with a 2.5% permanent income increase (in period 1 in the counterfactual economy) to achieve the same welfare as under the baseline policy. Here, we compute the population-weighted average CEV across all 750 cells.

**Implementation via Interpolation.** Given the complex nature of the expected utility function, computing the CEV directly is computationally intensive. Instead, we exploit the fact that we have already solved the model for 150 cells spanning the income distribution by the five health quintiles. For each income level, we have computed expected lifetime utility accounting for optimal savings, private ODI purchase decisions, and all subsequent consumption and transfer outcomes. Thus, our interpolation procedure works as follows:

1. For an individual with period 1 income $w$, compute their expected utility under the baseline policy: $U^* = EU^{Baseline}(w)$
2. Find two adjacent income grid points $w^{(j)}$ and $w^{(j+1)}$ such that: $EU^{\text{counterfactual}}(w^{(j)}) \leq U^* \leq EU^{\text{counterfactual}}(w^{(j+1)})$
3. Use interpolation to find the income level $w^*$ such that $EU^{counterfactual}(w^*) = U^*$
4. Compute the CEV: $CEV = \frac{w^*-w}{w} \times 100$

With 150 grid points, adjacent points differ by a small amount, implying a small interpolation error. As shown in Table G1, we validate this approach by showing that CEV calculations are nearly identical when using alternative interpolation methods, e.g., linear interpolation or shape-preserving piecewise cubic interpolation.[19]

**Table G1:** Robustness of CEV Calculations Across Interpolation Methods

| | (1) **Post-Reform** | (2) **↑ 10% WDI** | (3) **↑ 25% WDI** | (4) **↓ 10% cash transfer** | (5) **↓ 50% variable admin** | (6) **↓ 50% fixed admin** | (7) **Full Info** |
|---|---|---|---|---|---|---|---|
| Spline | 0.4394 | −0.1129 | −0.3505 | 0.3551 | −0.0528 | −0.0540 | 0.1081 |
| Linear | 0.4572 | −0.1107 | −0.3290 | 0.3747 | −0.0492 | −0.0552 | 0.1090 |
| Pchip | 0.4493 | −0.1134 | −0.3359 | 0.3686 | −0.0507 | −0.0569 | 0.1077 |
| Makima | 0.4519 | −0.1139 | −0.3355 | 0.3691 | −0.0506 | −0.0553 | 0.1077 |

*Note*: This table reports population-weighted average consumption-equivalent variation (CEV) in percentage points across policy counterfactuals using different interpolation methods. The four interpolation methods are spline interpolation using not-a-knot end conditions, linear, shape-preserving piecewise cubic interpolation (Pchip), and modified Akima cubic Hermite interpolation (Makima). Column (1) reports the percentage income increase required in the post-reform economy for individuals to be equally well-off as in the pre-reform economy. Columns (2) to (7) report the percentage income increase required in the counterfactual economy for individuals to be equally well-off as in the post-reform economy (Column (1)).

[19] To avoid imposing functional form assumptions beyond the grid boundaries, we employ linear extrapolation at the endpoints regardless of the interpolation method used for interior points. Since at most 0.8% of all points fall outside the grid boundaries, this conservative approach has a negligible impact on overall accuracy while maintaining methodological rigor.

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