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

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Abstract

This paper comprehensively evaluates a fundamental reform of the public Disability Insurance (DI) system in Germany. Effective 2001, cohorts born after 1960 are no longer eligible for "occupational DI." Occupational DI (ODI) implies benefit eligibility if employees are no longer able to work in their previous occupation. For the affected notch cohorts, the new general DI eligibility rules require that their reduced work capacity must prevent them from working in *any* job. Using administrative statutory pension insurance data, we first show that the reform significantly reduced the inflow of new DI beneficiaries by 20% for males and 10% for females. Next, we validate these findings using representative SOEP household panel data comprised of the entire underlying population and not just DI inflows. Moreover, at least at the population level, we do not find a significant overall increase in the likelihood to work full-time among notch cohorts, but some evidence for reduced subjective well-being. Next, using representative data on old age saving motives and health, we find no evidence that the notch cohorts purchased individual *private* ODI policies at higher rates to compensate for the reduced generosity in the public DI system. However, we do find evidence that a series of structural public pension reforms substantially increased demand for private old-age insurance among younger people in general. As German private ODI policies are individually underwritten and not guaranteed issue, we find strong selection based on observables—and also what are typically unobservables—into the relatively big private ODI market in Germany. While around 40% of younger cohorts have private ODI coverage, sick individuals are significantly less likely to be covered, as are individuals who expect to die young due to bad health and an unhealthy lifestyle. Those who believe it would be crucial to save for old age and unexpected life events purchase private ODI at much higher rates. By contrast, liquidity constrained households purchase private ODI at significantly lower rates. Finally, we categorize uninsured ODI households into three groups: (i) a third who are close to retirement and thus have low demand for private coverage despite being relatively healthy and wealthy, (ii) a quarter who are young but sick and low-income and thus cannot purchase private policies, and (iii) the remaining group of middle-aged households.

Keywords: occupational disability insurance, social safety net, private disability insurance, advantageous selection, private information, labor supply, well-being, health shocks, consumption, redistribution,

JEL Codes: H53, H55, I10, I14, I18, J14, J21, J26

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1. Introduction

For decades, the question of how to design "optimal" social insurance systems has been at the core of economic research (Chetty and Saez 2010; Chetty and Finkelstein 2013; Luttmer and Samwick; 2018, Goodman-Bacon 2018a; Cabral et al. 2019). Countries around the world have organized their social insurance and safety net systems differently; however, three integral parts of public social insurance—public unemployment insurance (Lalive et al. 2015; Hendren 2017), Workers' Compensation (Powell and Seabury 2018) and public disability insurance (Koning and Lindeboom 2015; Autor et al. 2016)—exist in basically every OECD country (OECD 2010; Pichler and Ziebarth 2020). What's more, compared to health insurance systems, their design and structure are similar across countries. Consequently, experiences from one OECD country might hold important lessons for others (Burkhauser et al. 2016).

In the United States, public disability insurance (DI) is one of the few relatively generous federal social safety net programs. As a result of rising beneficiary rates and spending, researchers and policymakers have analyzed and discussed the implications for labor supply, beneficiary health and well-being, multi-generational "welfare" cultures as well as household income, consumption and poverty (Dahl et al. 2014; Gelber et al. 2019; Autor et al. 2019). Using quasi-random case worker assignment, studies inside and outside the United States conclude that employment rates among marginally rejected applicants are 10 to 30 percentage points higher compared to marginally accepted applicants (Bound 1989; Chen and van der Klaauw 2008; von Wachter et al. 2011, Maestas et al. 2013; French and Song 2014, Kostøl and Mogstad 2014). Among economists, there is also consensus that the generosity of the public DI system and the stringency of the health screening process are major determinants of the inflow of cases (Autor and Duggan, 2003; de Jong et al., 2011). Finally, it is a stylized fact that receiving DI benefits is usually an absorbing state, which is why reform debates often surround the question of *how to prevent* DI take-up in the first place; for example, Burkhauser and Daly (2012) propose to incentivize employers to accommodate workers after a health shock.

In this paper, we comprehensively study a fundamental reform to the public DI system in Germany. We contribute to the literature by assessing a range of important outcomes simultaneously to paint a rich picture of how the reform not only altered public DI inflows and the private individual DI market, but also how it affected individuals' well-being, income, consumption and labor supply decisions after a health shock. We holistically study this reform in the context of the German welfare state which is known to be generous, but whose public DI system has become significantly less generous after the fundamental reform of 2001. Specifically, the reform significantly altered the eligibility criteria for people born in 1961 or after. Before 2001 and for those born before 1961, the relevant eligibility criterion was the work capacity *in the previous occupation* (or a comparable occupation in terms of income and social standing); that is, "Occupational Disability Insurance (ODI)." After the reform, the relevant eligibility criterion became the capacity to work in *any type of job*, that is, "Work Disability Insurance (WDI)." Besides Burkhauser et al. (2016) and Börsch-Supan et al. (2021), ours is one of the first papers to evaluate this reform and study how it affected what we call the "German notch cohorts" using primarily standard Regression Discontinuity (RD) approaches.

In a first step, using administrative data from the German Statutory Pension Insurance (*Deutsche Rentenversicherung*, DRV), we show that the reform significantly reduced the inflow of new DI beneficiaries, by 20% among males and 10% among females. We also study alternative social insurance routes for people who were no longer eligible for public ODI. Using representative household panel data from the German Socio-Economic Panel Study (SOEP), we confirm the significant decrease in the public DI recipiency rate using the universe of the underlying population, not just the select sample of DI inflows. However, we find no evidence for a general increase in the likelihood to work full-time. Next, we characterize households at risk of occupational or general work disability after a health shock following approaches similar to Burkhauser and Schroeder (2007) as well as Meyer and Mock (2019). Households whose household head occurs a health shock that threatens their ability to earn market incomes are significantly less likely to be college educated and have lower-incomes already prior to their health shock. After such a health shock, their likelihood to receive a disability pension and exit the labor force increase by almost 50%, whereas equivalized total income decreases by only 4% and equivalized post-tax income by 2.5% as a result of the relatively generous German welfare state and intrahousehold risk sharing (cf. Ortigueira and Siassi 2013).

Next, we use another set of high-quality data, accumulated from the accounting system of all German insurers. These high-quality data come from the German Association of Insurers (*Gesamtverband der Deutschen Versicherungswirtschaft,* GDV) that maintains a census database on all flows and stocks of newly signed insurance policies of all active German insurers. Comparing the dynamics of new private ODI policies sold per year relative to other insurance policies sold, for example life and pension insurance, we find clear evidence of a significant increase in the demand for private insurance around the time of the reform. However, not only did the demand for private ODI increase, but also demand of other old-age policies—and these increases are rather linked to a *series* of social safety net reforms at the time than to a specific single reform.

This finding of a *general* uptick in private old-age insurance demand is consistent with our subsequent findings using representative panel data from the SAVE (*Sparen und AltersVorsorgE in Deutschland*, "Saving for Old Age in Germany"*)* survey and the years 2001 to 2010. We use these data to characterize households who hold private ODI and study market demand in more detail. In particular, SAVE survey offers a fascinating set of questions related to risk aversion, expectations about life expectancy, health in general and relative to peers as well as saving motives and attitudes towards self-insurance. It helps us to identify and characterize households with a latent demand for private ODI, those who hold private ODI, and others whose demand for private ODI is low.

We also provide clear empirical evidence that the German private ODI market is advantageously selected (Fang et al. 2008; Einav and Finkelstein 2011, Soika 2018). On this nongroup market, no guaranteed issue exists and premiums are risk rated. We show that sick individuals are significantly less likely to purchase private ODI policies, whereas married individuals and the self-employed are much more likely to hold such policies. Using unique data on what is usually considered private policyholder information, we find that individuals who believe that they will die younger than people in their gender-age group are significantly less likely to purchase policies, as are liquidity constrained households and those who find it unimportant to save for unexpected life events. Our empirical analysis suggests that about half of all uninsured who show latent private ODI demand effectively cannot purchase private ODI as they are either denied coverage or offered prohibitively high premiums. This finding is consistent with a 2008 administrative survey among WDI beneficiaries who were asked about the reasons for not having private ODI (Märtin et al. 2012; 2014).

In the final part of the paper, we use two pre and two post-reform waves of a rich representative consumer expenditure survey to estimate the reform impact on household consumption among the notch cohorts, again using a standard RD design. Then we feed our various parameter estimates into the standard Baily-Chetty optimal social insurance framework following Meyer and Mock (2019). By doing this, we assess—under different assumptions and conditions—whether the equilibrium outcomes, after fundamentally reforming the public DI system, suggest that benefit generosity in the new system is too low.

2. Literature

This paper contributes to rich and important strands of the economic literature. In addition to the seminal papers on DI and optimal social insurance cited above, we contribute to the literature that analyzes the interplay between public and private insurance markets and their regulation. This includes, but is not limited to, research analyzing health insurance markets. We also contribute to research that studies the intersection of coexisting social insurance systems and spillover effects between those (Borghans et al. 2014; Lalive et al. 2015; Leung and O'Leary 2020; Ahammer et al. 2020).

Regarding the interdependencies of public and private health insurance markets, rich strands of theoretical and empirical economic research have studied whether offering public health insurance crowds-out the demand for private health insurance. Most studies investigate *how large* such potential crowd-out effects really are, and how they can be empirically identified. Theoretically, crowd-out of private coverage may occur in the moment when government-provided or government-subsidized coverage is introduced or coexists with private, unsubsidized, coverage options (Pauly 1990).

In the U.S., a first set of empirical papers have studied whether expanding the state-level Medicaid program for low-income pregnant women and children reduced private insurance coverage in the 1980s. Cutler and Gruber (1996) find large crowd-out effects implying that the number of privately insured children decreases by 30 for every 100 children who are newly eligible for Medicaid. However, a series of follow-up studies using different datasets, methods, and crowd-out definitions—and evaluating reforms in late 80s or early 90s—provide very mixed and inconsistent findings: from little to no crowd-out effects (Card and Shore-Sheppard 2004) to quite substantial estimates, see Gruber and Simon (2008) for a literature review. One conclusion is that heterogeneity in crowd-out across eligibility groups and income categories certainly exists (Koch 2013, Garthwaite et al. 2014, Dague et al. 2017).

A second set of papers have studied related crowd-out questions. For example, Glied and Stabile (2001) provide evidence that anti-crowd-out provisions like the 1982 Medicare Secondary Payer Provision triggered low compliance rates and may have led to unintended employer and employee behavior. Lo Sasso and Meyer (2006) find evidence that the existence of uncompensated care by hospitals may crowd-out demand for health insurance; a hypothesis that is confirmed in calibrated simulations by Qin and Liu (2013). Wagner (2015) finds large crowd-out effects of up to 100% when studying Medicaid expansions for the disabled. Other papers study intra-household crowd-out spillovers: Koch (2015) finds that children's public insurance can crowd-out private insurance of parents. And Witman (2015) exploits age gaps between spouses and shows that the Medicare eligibility of an older spouse can crowd-out private coverage of a younger partner. In one of the very few studies on crowding-in, Clemens (2015) shows that Medicaid's coverage of unhealthy adults actually increases coverage rates in private community-rated markets by improving the quality of the remaining risk pool.

A last set of papers investigates whether Medicaid would crowd out the demand for private long-term care insurance and reduce private savings (Sloan and Norton 1997; Gruber and Yelowitz 1999, Brown et al. 2007; Brown and Finkelstein 2008). Moreover, employer-provided retiree health insurance could crowd-out household wealth (Clark and Mitchell 2014) and the existence of family members (who could provide informal care) could crowd-out the demand for private long-term care (Pauly 1990).

Ours is one of the first papers to test for the link between private and public disability insurance markets, and whether reduced public DI can *crowd-in* private DI take-up. For example, Autor et al. (2014) study the private DI group market in the United States and also discuss the relation to the public DI market but do not formally test any crowding-in or out. Burkhauser et al. (2016) describe the German market in comparison to other countries and suggest that the 2001 reform may have induced private market crowding-in.

3. The German Disability Insurance System

Social Insurance in Germany

Germany, like most European nations, has a generous social safety net consisting of public Unemployment Insurance (UI), Workers' Compensation (WC), Health Insurance (HI) and Long-Term Care (LTC) insurance (cf. Schmieder et al. 2016; Bauernschuster et al. 2020). In addition, among employees, eligibility for sick and medical leave is universal (Ziebarth and Karlsson, 2010, 2014; Ziebarth 2013). Moreover, Germany runs a Statutory Pension Insurance (SPI) program (Eibich 2012, Geyer, 2021), and also a universal needs-based cash transfer program that provides a guaranteed social minimum income floor to all its citizens called *Unemployment Insurance II*[1] for citizens who are generally able to work and part of the labor force (Konle-Seidl 2012; Dustman et al. 2014).[2] The programs are funded through a mix of contribution rates for UI, WC, HI, LTC and SPI as well as employer mandates for paid sick leave and general taxes for the social minimum income floor. See Eichhorst et al. (2008), Ziebarth (2018) and McVicar et al. (2021) for more detailed overviews.

History of Public Disability Insurance in Germany: 1970 to 2001

Germany's public DI program is part of SPI. It provides benefits for both partially and totally disabled workers, who have paid contributions during their work life. Employers and employees are each subject to a payroll tax—9.3 percent—of their monthly gross wage up to the social insurance contribution ceiling.

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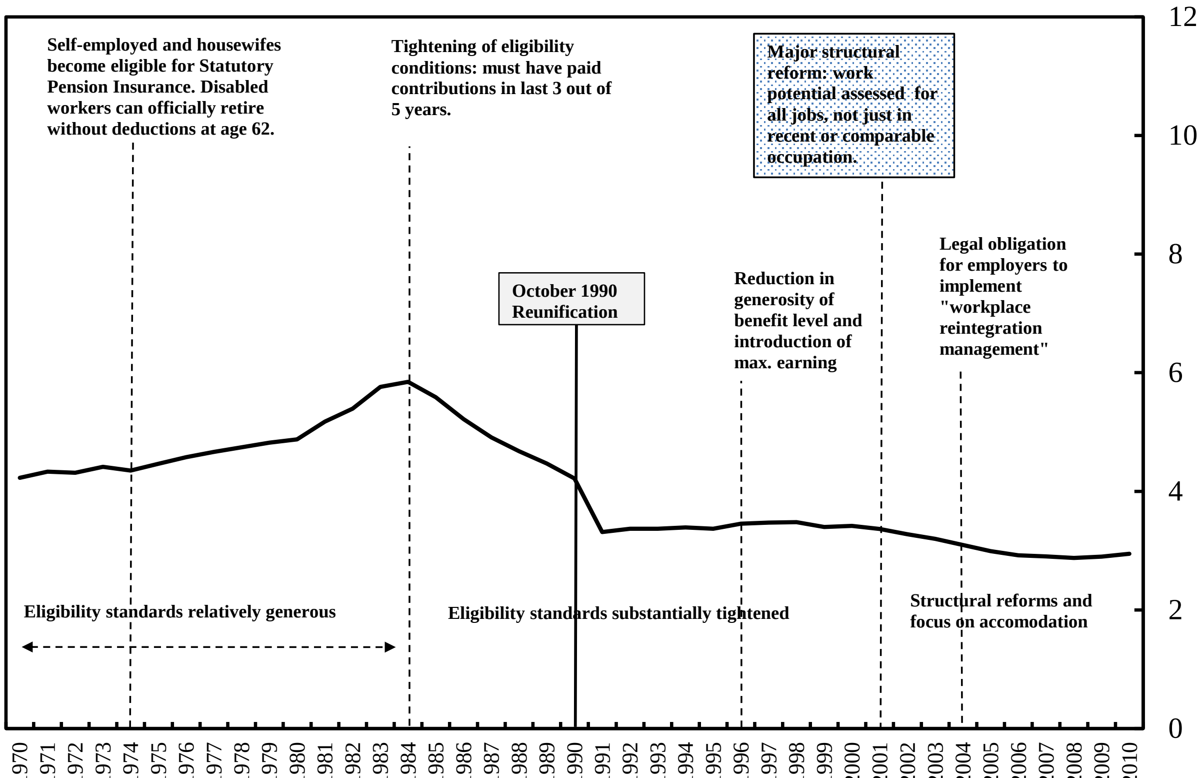
[1] The 2004 reforms created the *Arbeitslosengeld II* program (*Sozialgesetzbuch II*, "Social Code Book II"). For more information about the reforms see, e.g., Eichhorst et al. 2008, Konle-Seidl 2012.

[2] People are considered to be "able to work" if they are able to work at least 3 hours per day. A relatively small share of people receive *Sozialhilfe (Hilfe zum Lebensunterhalt)* ("Social Assistance Benefits") of a similar amount but have no job search requirements and are not considered to be in the labor force (§§27-40 SGB XII). These beneficiaries are typically "long-term unemployed" and classified as *temporarily* not able to work 3 hours per day.

Figure 1 shows the development of Germany's public DI caseload from the 1970s to 2014 along with selected reforms. In the early 1970s, Germany had high disability recipiency rates compared to other OECD countries such as the US, the UK, Sweden, the Netherlands or Australia (Burkhauser et al. 2016). What potentially contributed to this high rate was a change in eligibility rules in 1969. In that year, the highest German Social Court ruled that partially disabled workers are eligible to full benefits if they are unable to find a job (Burkhauser and Hirvonen, 1989; Burkhauser 1991). This rule has not changed since then. At the time, about half of all pension entries occurred through the DI scheme. A major generous reform in 1972 introduced new possibilities to receive retirement benefits earlier than the statutory retirement age of 65 without actuarial deductions. At first, this reduced DI entries but, as seen in Figure 1, DI enrollment rates rose significantly, peaking at 5.8 percent of the workforce in 1984.



Source: Deutsche Rentenversicherung (2014), Statistisches Bundesamt (2014)

Figure 1: DI Recipiency Rates as a Share of the Working Population and Reforms
(adapted from Burkhauser et al. 2016)

In 1982, the newly elected center-right government restricted eligibility to workers who had paid SPI contribution rates over the past three out of five years. As many housewives (or househusbands) did not meet these criteria, the strong decline in DI recipiency rates between 1984 and 1990 were linked to restricting access for mostly women working outside the formal labor market. The inflow of new female recipients fell sharply from about 173,000 in 1984 to 67,000 in 1986 as a result of the reform (see Robert Koch Institute (2006) as well as Börsch-Supan and Jürges (2012) for a more detailed discussion).[3]

In the 1990s, in reunified Germany, additional reforms were launched. Specifically, caps on the earnings of beneficiaries were introduced in 1996. A strong reduction in the inflow of new beneficiaries contributed to the decline in the overall DI recipiency rates over the rest of the decade, as evidenced in Figure 1.[4] Already at that time, the center-right government send strong signals and prepared reforms packages to cut the generosity of the old-age pension system while sending a message that private savings and insurance plans would be essential in the future.

The Fundamental Public Disability Insurance Reform of 2001

The most fundamental of all DI reforms became effective in 2001. It systematically changed the federal DI system and the Social Code Books that include the legal specifics.[5] Until 2001, the system included two disability pension schemes: occupational and general disability. General disability status was granted if a person could not earn more than a minimum income. Occupational disability required that one is "*unable to work in the occupation in which one was trained—effectively in the last job or a comparable job in terms of the skills it required, the wages it paid and its prestige*." General disability gave access to full disability benefits, and occupational disability benefits were two thirds of full DI benefits. A central element of the 2001 reform was the

[3] At the same time, the waiting period for old-age pensions was reduced from 15 to 5 years. Therefore, more women with short labor market careers became eligible to retirement benefits at age 65.
[4] Note that the figures reflect the stock of all beneficiaries. As such, even large declines in the inflow of new beneficiaries only gradually translate into overall DI rate declines.

[5] An entirely new Social Code Book IX (*Sozialgesetzbuch IX*) was introduced that regulated the Rehabilitation and Participation in Social Life (*Rehabilitation und Teilhabe Behinderter Menschen*) for disabled people in Germany. Before 2001, most of these regulations were included in the *Schwerbehindertengesetz.*

abolishment of this occupational status protection. Instead, DI benefit eligibility depends now on the ability to work in *any* occupation. However, a grandfathering clause for people born before January 2, 1961, was agreed upon such that, effective January 1, 2001 the previous Occupation Disability Insurance (ODI) was converted into a Work Disability Insurance (WDI) for cohorts born after December 31, 1960.

In the new WDI system, caseworkers evaluate the work capacity based on medical diagnoses. Instead of a potential earnings threshold, after 2001, work capacity is measured by working hours. If the general work capacity lies between 3 and 6 hours per day, then partial WDI is granted (50% of full benefits); and if the general work capacity is less than 3 hours per day, full WDI is granted (Deutsche Rentenversicherung 2020b). If people are eligible to ODI under the grandfathering clause, they receive partial WDI. Note that this is a lower benefit level (50%) than the 66% in the previous regime. As mentioned above, if there is no suitable part-time job offer, applicants are eligible for full DI benefits.[6] Moreover, after the 2001 reform, WDI benefits are generally granted temporarily for three years. After nine years, it has to be converted into a permanent benefit. If health and working abilities are not expected to improve, a permanent pension can be granted earlier.[7]

The main eligibility criteria did not change in the course of the 2001 reform. These require applicants to have paid social security contributions in the last three out of five years, and a general waiting period of five years exists as well. When granted, beneficiaries receive benefits as a type of "early retirement pension" with the associated actuarial reductions applied. Benefits are mainly based on individual earnings histories and not adjusted for family composition, income or assets. Benefits are calculated the same way as old-age pensions assuming that the person earns his/her individual average earnings over the rest of the career until age 60.

Actuarial deductions of 3.6 percent per annum are applied if the individual retires before 60. However, deductions are capped at 10.3 percent, that is, all individuals who enter DI benefits before 57 are affected by maximum deductions. Average benefits are relatively low and the 2001 reform lead to a further and long-lasting

[6] This is usually assumed to be the case if the DI recipient cannot find work within a year. The share of partial DI benefits converted to full DI benefits lies between 12 and 16% of all pension entries in the years between 2002 and 2019 (Deutsche Rentenversicherung 2020c), earlier data are not available.
[7] In 2019, about half of all new WDI beneficiaries received temporary benefits (Deutsche Rentenversicherung 2020c).

decline in benefits. For example, average benefits of new recipients were €713 in 2000, and—despite regular pension increases—declined by 14% to €640 in 2010. In this sense, the fundamentally restructured WDI system is truly a last resort program for those whose health impairments keep them from doing any job in the economy for less than three hours per day.[8]

In this paper, we will comprehensively evaluate the effects of the 2001 reform on a range of outcomes using several different datasets. In principle, our identification strategy will compare those cohorts whose public DI was downgraded from ODI to WDP effective 2001 to cohorts who could still enjoy the much more generous ODI benefits after 2001. As seen in Figure 1, when using aggregated administrative data, the stock of beneficiaries does not seem to decrease sharply due to the reform. However, as we will see, a more refined analysis will unravel effects on various outcome margins.

Moreover, further reforms in 2004 continued to reduce the inflow of new beneficiaries. However, the attention shifted away from tightening WDP eligibility requirements towards promoting worker accommodation on the job. Specifically, the reforms mandated employers to provide "workplace reintegration management" (*Betriebliches Eingliederungsmanagement*, §84 SGB IX). Indeed, the law requires that when impaired workers exhaust their short-term sickness benefits (six weeks) and is being considered for longer-term sickness benefits (Ziebarth 2013), employers must coordinate a reintegration plan considering input from the sick-listed employee, work disability experts (*Integrationsämter*), the work council (*Betriebsrat*), and the workplace physician

[8] Also codified in Social Code Book IX (SGB IX) is a coexisting Disability Classification System (DCS). This DCS system assigns citizens with health impairments a handicap rating by medically-based impairment categories. Only permanent health impairments lead to a classification. For example, a mild form of Parkinson disease without imbalance issues but "mild motion disorders" yields a handicapped degree of 30-40% (BMAS, 2009). Handicapped rating of 50% and above imply "severely handicapped" and "social disadvantage compensations" (*Nachteilsausgleich*) such as special income tax deductions, the ability to retire two years earlier without deductions, or parking lots for people in wheelchairs. Effectively, all public DI beneficiaries would be eligible for these benefits without any loss of benefits as the DI system is not means tested and as the two systems are independent. DI benefits are solely based on work capacity; being handicapped does not automatically imply eligibility for DI benefits. About 1.1 million severely handicapped people work full time in Germany (Bundesagentur für Arbeit 2020). This is, at least partly, the result of a quota system according to which the workforce of employers with more than 19 full-time employees has to be composed of at least 5% severely handicapped workers. Otherwise, employers have to pay a monthly penalty (*Ausgleichsabgabe*) of €290 per unoccupied workplace. Using data from Austria with a similar system, Lalive et al. (2013) show that such a quota system significantly increases employment of handicapped people.

(*Betriebsärzte*). The general principle as codified in its Social Code Book IX is "*Rehabilitation before Pension*" emphasizing prevention and measures to maintain life-long working capacity. That is, the idea behind the 2004 reform is to overcome temporary disability and to prevent future deteriorations in work capacity. However, this reform is not in the focus of this paper and affected all birth cohorts equally. It likely had a gradual impact on the decreasing stock of DI beneficiaries as seen in Figure 1.

Another point is worth mentioning in this context. The reduced DI benefit level increased the share of DI recipients who also apply for means-tested social assistance. In 2019, only 2.6% of all pensioners received social assistance; but among people with DI benefits, the rate was nearly 15% (Deutsche Rentenversicherung, 2020c). A series of studies using different data sources documents this increasing risk of poverty (Krause 2013, Märtin et al., 2012; 2014; Geyer, 2021). As a consequence, in a series of reforms after 2014, policymakers responded to this development and increased new DI benefit pensions by increasing the reference age for the calculation of DI benefits. As described, the reference age was 60 since 2001. Starting July 2014, it was increased to 62, then modestly in 2018, and in 2019, it increased to 65 years and 8 months. Now it equals the statutory retirement age and will increase to 67 years by 2031. Between 2018 and 2019, benefit levels of new DI recipients increased by about 10%. At the end of 2019, total WDP benefits per month were €1.2 billion, or 8% of total spending by the SPI (Deutsche Rentenversicherung, 2020a).[9]

Private Disability Insurance in Germany

Below we will estimate the impact of the 2001 reform on the demand for private ODI. The German private disability insurance market is overwhelmingly an individual market, not a group market like in the United States (cf. Autor et al. 2014). Similar to the long-term health insurance market in Germany (Atal et al. 2018, 2020), the private individual ODI market is individually underwritten and guaranteed issue does not exist. Private disability insurance follows private insurance law *(Versicherungsvertragsgesetz)*. It is based on a private contract

[9] The figure of €1.165 billion (14 billion per year) is based on 72,301 partial WDI beneficiaries with average monthly cash benefit of €553 and 1,415,295 full WDI beneficiaries with average benefits of €842 (Deutsche Rentenversicherung, 2020a). Note that WDP benefits are converted to an old-age pension of the same amount when reaching the statutory retirement age. Expenditures are calculated for WDI beneficiaries younger than the statutory retirement age.

between the insurer and the insured, which specifies the conditions for the insured risk individually. Premiums depend on age, medical diagnoses, and occupation. As a result, premiums can be high for high-risk occupations and applicants may be denied coverage. In 2012, 61 percent of employed men and 42 percent of employed women were covered by private disability insurance, which is almost always ODI coverage (Statistika, 2014). In 2015, according to the *German Association of Insurers* (GDV), the average pension from a private ODI was at €7,551 per year.[10] Data from 2018 show that about 50% all contracts include coverage of more than €10,000 per year.[11]

In contrast to the U.S. market, where private group DI usually include "offset clauses" that may reduce public Social Security Disability Insurance benefits dollar for dollar (Burkhauser and Daly 2011), in Germany, private and public DI benefits do not crowd-each other out.

4. Impact of 2001 Reform on Demand for Public and Private Disability Insurance

In a first step, we use administrative data to provide evidence on the impact of the 2001 reform on new DI inflows by birth cohort and year in a difference-in-differences (DD) framework. Next, we use representative household panel data from the German Socio-Economic Panel Study (SOEP) to validate the findings with the universe of the underlying cohort populations in a regression-discontinuity (RD) design. Moreover, using the SOEP, we assess the effects on labor supply, income and subjective well-being among notch cohorts. After that, we use a census of private insurance policies sold in Germany from 1981 to 2013 to document in a time series framework how aggregate demand for various policies has changed over time. After providing this first snapshot about the private insurance market, we use representative survey data from 2000 to 2010 to pinpoint possibly increased private demand by the notch cohorts. Further, using an unusually rich set of demand predictors, including health measures and individuals' subjective assessment of their life expectancy, we carve out actual and latent private insurance demand and categorize uninsured households. Recall that many individuals face

[10] his figure includes full ODI and supplemental contracts, see https://www.gdv.de/de/themen/news/75-prozent-aller-kunden-ohne-wenn-und-aber-aufgenommen-31438 (last accessed on March 25, 2021)

[11] https://www.gdv.de/de/themen/news/versicherte-renten-steigen-an-59814 (last accessed on March 25, 2021)

prohibitively high premiums or coverage denial in a risk-rated market without guaranteed issue. In the final part, we use representative household expenditure data to estimate the reform effects on household consumption and discuss possible welfare effects in a standard Baily-Chetty framework.

4.1 Public DI Inflows and Alternative Social Insurance Routes Using Administrative Data

In the first part, we use administrative data from the German Statutory Pension Insurance (*Deutsche Rentenversicherung*, DRV) on the inflow of public DI beneficiaries, separately by year, age and cohort.[12] We normalize the number of inflows by cohort population size in each year.[13] Using data from 1995 to 2018, we then study whether the 2001 reform reduced public DI inflow rates for the notch cohorts who experienced a downgrading of their policies from ODI to WDI. To recall, these were all cohorts born after 1960, that is, those who were 40 years of age on January 1, 2001. Consequently, we run the following Difference-in-Differences (DD) model

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

where y_{ct} denotes the share of new public DI recipients of cohort *c* in year *t*; D_c is a dummy that identifies notch cohorts, that is, people younger than 40 as of January 1, 2001; T_t is a post-reform indicator that turns on 1 after 2000, δ_t are year fixed effects and ρ_c are cohort fixed effects. e_c denotes the error terms, which we cluster at the cohort level. The main identification assumption implies that, absent the reform, the inflow of new public DI beneficiaries in the notch cohorts would have developed in the same manner than the non-notch cohorts.

To illustrate the main findings, Figure 2 plots an event study using equation (1) but replacing T_t with a series of year dummies where 2000 serves as baseline year (Goodman-Bacon 2018b). As seen, whereas the five pre-treatment years show no trending and relative inflow differences between treated and control cohorts are

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[12] For the years 1995 to 2009 we received aggregated data of all disability pension inflows by gender, age and region for the years 1992 through 2010 from the statistical office of the DRV (standardized Tables 201.00 Z_, 201.01 Z_, 201.02 Z_, 201.10 Z_, 201.11 Z_, 201.12 Z_, 201.20 Z_, 201.21 Z_, 201.22 Z_). Since 2010, data are available from the website of the statistical unit of DRV (https://statistik-rente.de/drv/).

[13] We use aggregate population data from the Genesis database of the Federal Statistical Office (*Statistisches Bundesamt*): https://www-genesis.destatis.de/genesis/online (last accessed March 20, 2021).

not significantly different from zero, we observe a sharp decline beginning in the first post-reform year 2001. This decline further accelerates in subsequent years, up to point estimates exceeding -0.2 percentage points, or about 35% relative to the pre-reform mean.[14] In 2011, two decades after the reform implementation, when the first notch cohorts turned 60, the inflow differential between the two groups stabilizes and remains highly significant at -0.2 percentage points.

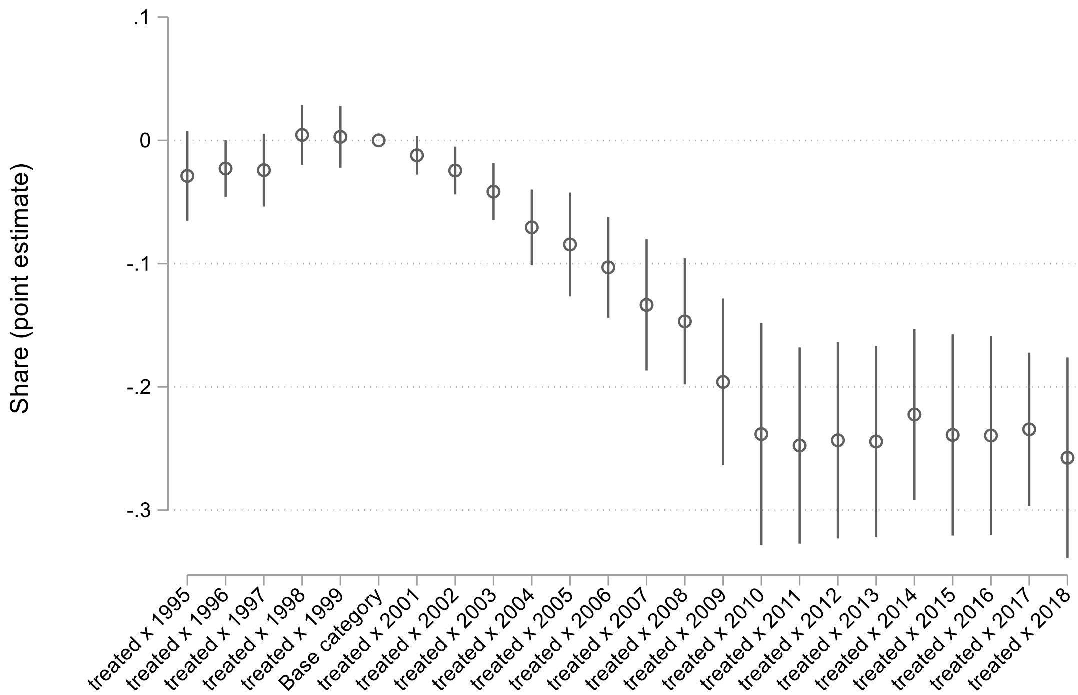

Figure 2: Effect of 2001 Reform on Public DI Inflows—Event Study All

Figure A1 (Appendix) shows the same event studies separately by gender. We observe the same reassuring stable pre-reform trends, followed by substantial inflow reductions among the notch cohorts. However, not surprisingly, the reform-induced decrease in inflows is substantially larger for males as their eligibility rates are higher due to a stronger labor market attachment. Men are more likely to fulfill the requirement of having three

[14] We define the pre-reform mean as the mean entry rate of untreated cohorts in this regression (1954-1960, between 32 and 58 years of age) which is 0.58%.

years with compulsory contributions over the last five years.[15] Moreover, men are more likely to work in physically demanding occupations and industry jobs and generally face a higher disability risk.

Table 1: 2001 Reform Effect on Public Disability Insurance Inflows using Pension Insurance Data

		(1)	(2)	(3)	(4)	(5)
All						
	$D_c \times T_t$	-0.0907***	-0.0907***	-0.0907***	-0.144***	-0.0514***
		(0.0293)	(0.0219)	(0.0184)	(0.00992)	(0.0105)
	T_t	0.364***	0.485***	0.485***	0.762***	0.774***
		(0.0199)	(0.0344)	(0.0289)	(0.0192)	(0.0204)
	D_c	-0.159***	-0.266***	-0.266***	-0.397***	-0.0782***
		(0.0255)	(0.0290)	(0.0243)	(0.0137)	(0.0101)
	East Germany			0.175***	0.174***	0.169***
				(0.00782)	(0.00416)	(0.00450)
	Female			-0.0411***	-0.0325***	-0.0284***
				(0.00782)	(0.00416)	(0.00450)
	N	1,300	1,300	1,300	1,164	388
	Control group mean	0.61	0.61	0.61	0.58	0.50
Men						
		-0.127***	-0.127***	-0.127***	-0.174***	-0.0649**
	$D_c \times T_t$	(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
Women						
		-0.0548**	-0.0548**	-0.0548**	-0.115***	-0.0378**
	$D_c \times T_t$	(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
	Age groups	20-64	20-64	20-64	32-58	32-58
	Cohorts	1954-1966	1954-1966	1954-1966	1954-1966	1959-1962

Notes: German Pension Insurance, administrative data on public DI inflows, 1995-2018. Each column in each panel is from one DD model as in equation (1). Panels for men and women control for D, T and East Germany but coefficients are omitted for readability. See main text for more details.

[15] In 1985, when this requirement became effective for the first time, disability pension inflows of women dropped by about 50% whereas male inflows only slightly decreased, see Figure 1 and Section 3 as well as Deutsche Rentenversicherung, 2020c).

Table 1 shows the parametric DD model equivalents, where the upper panel shows the results for the full sample, the middle panel shows the results for men, and the bottom panel shows the results for women. Each column in each panel stands for one separate DD mode like in equation (1). The findings in Table 1 are in line with the event study estimates. First, the estimates are robust to the inclusion of cohort and year fixed effects as well as controls for East Germany. The average decline in inflows for males translate into a 20% decrease relative to the mean of the control group. The decline for women is only half as large at 10%. However, when zooming-in and restricting the bandwidths of cohorts considered, e.g. to cohorts born in 1959 to 1962, the effect sizes decrease to -12.5% for males and 7.9% for females.

4.2. Labor Supply, Income and Well-Being Effects Using SOEP Data

In the next step, we extract data from the representative German Socio-Economic Household Panel (SOEP); see Goebel et al. (2019) for more details on the dataset. Specifically, we focus on the years 1995 to 2016 and on males below the age of 60. In addition, we focus on birth cohorts from 1940 to 1980. Table A1 shows the summary statistic, where we list our main outcome variables in the upper panel and the sociodemographics that we use as control variables in the lower panel.

That allows us to run the following standard Regression Discontinuity (RD) model:

$$y_{it} = \alpha + \beta D_i + \gamma_0 (1 - D_i) f(z_i - c) + \gamma_1 D_i f(z_i - c) T_t + X'_{it}\tau + \delta_t + \rho_s + e_{it} \qquad (2)$$

where y_{it} measures whether the household receives public DI benefits and D_i is one if the respondent belongs to the notch cohorts. The cohort measure z_i enters the empirical model in difference to the reform cutoff c, which is 1961. In our baseline specifications, we include a linear and cubic trend in the running variable $f(z_i - c) = z_i - c$ which allows for different slopes before and after the cutoff. All regressions include year (δ_t) and state (ρ_s) fixed effects. X'_{it} is a rich set of socio-demographic, educational and job-related control variables as listed in the summary statistic in Table A1. For example, the average age is 42, 65% are married and between 0 and 11 children belong to the household, which have on average 2.2 members. About 22% finished the highest educational track in Germany and 4% are part-time employed; 35% are white-collar employees.

The standard errors e_{it} are clustered at the level of the running variable, which is the cohort level. The main RD identification assumption implies that no other factor that would affect public DI caseloads trends discontinuously at the birth-cohort level. We are not aware of another reform or factor that could invalidate this assumption. Note that we are now using the universe of the underlying population of interest and not just select DI inflows as with the administrative data. Thus, it allows us to study the impact of the 2001 reform on public DI receipt using an RD framework as in equation (2).

The SOEP Group provides a time-consistent longitudinal binary variable that indicates whether individuals receive an old-age pension due to work disability. We call this variable *Public DI I*. Moreover, they provide a second generated variable indicating the annual income stream from old age, disability or civil servant pensions. Using only respondents with a positive pension amount who do not receive a civil servant, a veteran's, a miners' or a farmers' pension, we create a second binary indicator to flag recipients of public DI pensions, *Public DI II*.

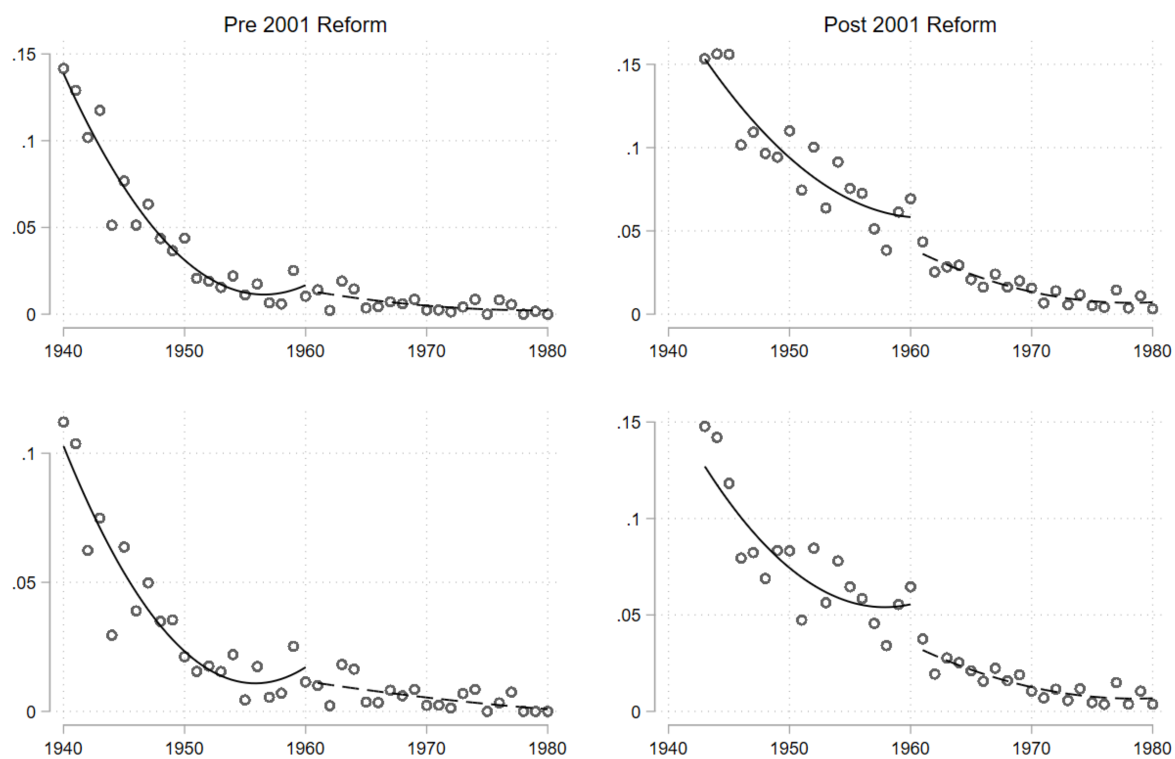

Figure 3: Public Disability Insurance Receipt by Birthyear

Figure 3 plots Public DI recipiency rates by birth cohorts. The first row shows unconditional scatters along with local polynomial smoothing cubic plots for *Public DI I*, and the second row for *Public DI II*[16]; the first column shows the results for the pre-2001 years and the second column for the post-2001 years.

Table 2: Impact of 2001 Reform on Public DI Recipiency Rates: 2001-2016

	(1) Public DI I	(2) Public DI II	(3) Non-Married	(4) Single Household
D	-0.018*	-0.020**	-0.010	-0.033*
	(0.0100)	(0.0095)	(0.0134)	(0.0191)
(1-D)*yob	0.018***	0.015***	0.013***	0.021***
	(0.0027)	(0.0024)	(0.0034)	(0.0050)
D*yob	0.014***	0.011***	0.013***	0.017***
	(0.0020)	(0.0015)	(0.0023)	(0.0025)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
R-squared	0.044	0.034	0.056	0.049
D	-0.019**	-0.021***	-0.038***	-0.041***
	(0.0079)	(0.0077)	(0.0131)	(0.0128)
(1-D)*yob	0.005	0.004	0.011***	0.011**
	(0.0028)	(0.0025)	(0.0038)	(0.0042)
D*yob	0.003	0.002	0.010**	0.010***
	(0.0022)	(0.0018)	(0.0037)	(0.0029)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
socio-dems	X	X	X	X
education	X	X	X	X
R-squared	0.194	0.162	0.174	0.235
Control cohort mean	0.0379	0.0323	0.0414	0.0563
Observations	87,472	87,472	29,109	31,889

Notes: SOEP v.33 -- 95% sample. Each column in each panel is from one RD model as in equation (2), enriched with cubic terms. The first two columns use thef ull sample, whereas the third column selects on non-married individuals and the final column on single households. The dependent variable is Public DI I, except in the second column where it is Public DI II. See main text for more details.

[16] Here we use cubic plots as they fit the raw data better. In Figure 7 below, we use linear plots. The linear plots for Figure 7 and cubic plots for Figure 3 are available upon request.

The visual evidence corroborates the findings from our administrative data above: we see a clear discontinuous decrease in the probability to receive a public DI pension for the notch cohorts in post-reform years. By contrast, for pre-reform years, being born after 1960 shows no obvious discontinuity in the likelihood to receive a public DI pension.

Table A2 shows the regression results from a model as in equation (2) for pre-reform years. Column (1) shows the findings for the full sample, column (2) uses the alternative *Public DI II* measure, column (3) focuses on unmarried respondents, and column (4) on single households. The upper panel uses RD models without control variables other than age, gender, state and year fixed effects, and the lower panel adds the socio-demographic, educational and labor market controls listed in the SOEP summary statistic in Table A1. As seen, the relevant point estimates are small and only one out of eight is marginally significant. For example, the point estimates in column (1) are -0.8 percentage points (upper panel) and -0.5 percentage points (lower panel).

Table 2 follows the same set-up but shows the results for the post-reform period from 2001 to 2016. Here, we find statistically significant results for seven out of eight models; all eight estimates carry consistently negative point estimates, in line with the right column of Figure 3. For example, the point estimates in column (1) are -1.8 percentage points (upper panel) and -1.9 percentage points (lower panel). Relative to the mean recipiency rate of the non-treated cohorts, 7.5%, these estimates translate into decreases of 24% (upper panel) and 25% (lower panel) and are very consistent with the results in Table 1.

Figure 4 repeats the exercise but focuses on the outcomes *Non-Employed*, *Full-Time Employed*, *Total Annual Income (equivalized)* and *Subjective Well-Being*. Here, we focus on the post-treatment period from 2001 to 2016. The first visual inspection does not yield much evidence for any discernable discontinuous jump, with the exception of subjective well-being maybe. The parametric findings in Table 3 confirm this conjecture. Here we always use the full sample but otherwise the setup of the table is similar to above.[17] As seen, the coefficients

[17] Results for the subsamples as in Table 3 are robust and available upon request.

in the first two columns are consistently very small and insignificant. The estimate in column (3) regarding the

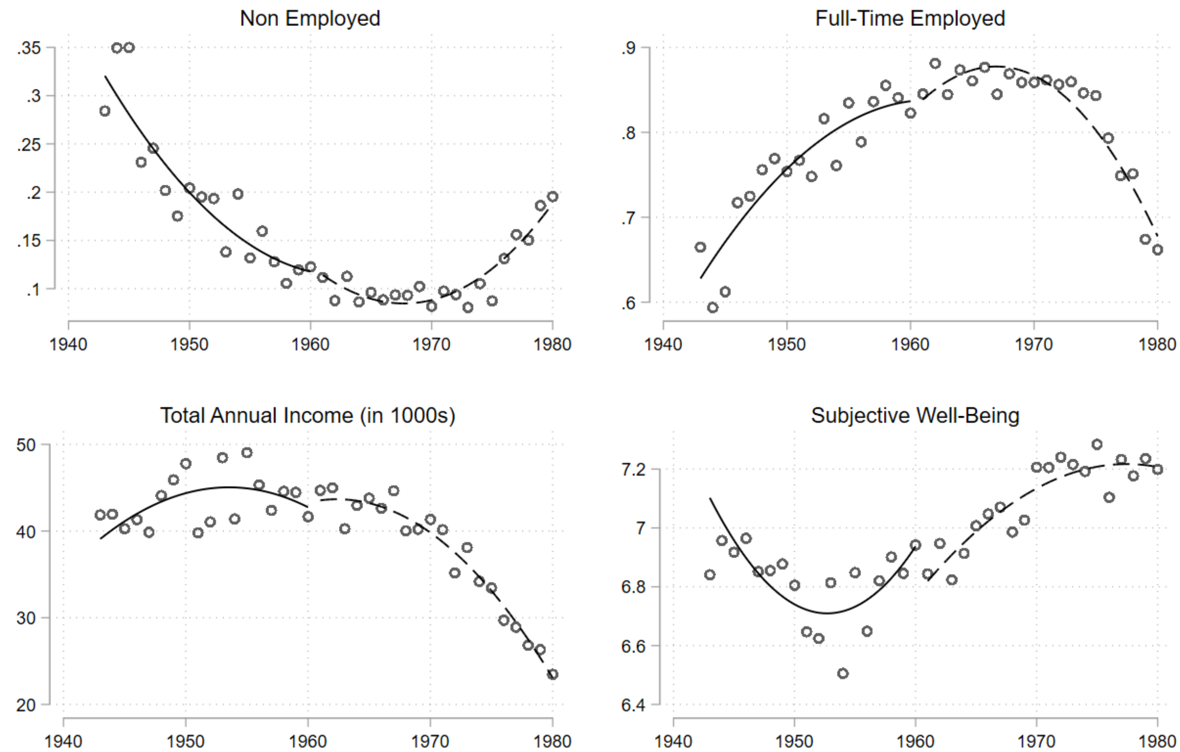

Figure 4: Labor Supply, Income and Subjective Well-Being by Birthyear

total annual equivalized income is insignificant in the upper panel and marginally significant and positive in the lower panel but overall not very precise. One could have hypothesized that incomes would decrease significantly after the 2001 reform. However, keep in mind that Germany runs a generous welfare with many alternative social insurance pathways other than public DI. As we learned from Section XX., administrative data show a significant increase in the number of XXX. Thus, given the overall relatively small reform effect of a -1.9 percentage point reduction in public disability pension receipt at the population level (column 1, Table 2), it is not surprising to find no evidence for negative income effects among notch cohorts in Figure 4 and Table 3. However, column (4) of Table 3 shows a potentially negative impact on subjective well-being among the notch cohorts. Below, we will further investigate this possible negative impact on utility by studying the reform effect on consumption among notch cohorts.

Table 3: Impact of 2001 Reform on Labor Supply, Income and Subjective Well-Being

	(1) Non Employed	(2) Full-Time Employed	(3) Individual Total Income	(4) Subjective Well-Being
D	-0.005	0.000	2.294	-0.116**
	(0.0114)	(0.0164)	(1.5386)	(0.0476)
(1-D)*yob	0.001	-0.001	-0.184	0.037**
	(0.0035)	(0.0040)	(0.1143)	(0.0165)
D*yob	-0.009***	0.014***	-1.068***	0.039***
	(0.0018)	(0.0028)	(0.0917)	(0.0093)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
Control cohort mean	0.1593	0.7987	43.0308	6.812
N	87,472	87,472	87,472	87,472
R-squared	0.031	0.039	0.077	0.036
D	-0.005	-0.000	2.342*	-0.102*
	(0.0103)	(0.0146)	(1.3806)	(0.0519)
(1-D)*yob	0.043***	-0.064***	-3.027***	0.067***
	(0.0057)	(0.0086)	(0.3008)	(0.0237)
D*yob	0.036***	-0.052***	-2.275***	0.057***
	(0.0046)	(0.0077)	(0.2688)	(0.0162)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
socio-dems	X	X	X	X
education	X	X	X	X
Control cohort mean	0.1593	0.7987	43.0308	6.812
N	87,472	87,472	87,472	87,472
R-squared	0.079	0.096	0.177	0.069

Notes: SOEP v.33 -- 95% sample. Each column in each panel is from one RD model as in equation (2), enriched with cubic terms. The dependent variables are listed in the column header. See main text for more details.

4.3. Labor Supply, Income and Well-Being after a Health Shock Using SOEP Data

We continue to use the representative SOEP for another exercise. Namely, to investigate how health shocks that threaten people's work capacity affect the likelihood to receive public DI benefits, labor supply, income and well-being. For that purpose, we follow Burkhauser and Schroeder (2007) and use the SOEP *work limitation due to health* measure to flag people with health shocks that threatens their work capacity. In particular, we indicate that respondents had a work capacity threatening health shock if they responded "*Yes, strongly limited*" (and did not in the previous wave) to the question of whether their health would limit their capacity to work. Unfortunately, this question was only asked in post-reform survey years 2011, 2012, 2013, and 2015.

Then we regress our outcome variables of interest on the lagged health shock regressor (as it takes time to apply for and get approved DI benefits).[18] The results are in Table A3, where the column headers indicate the outcome variable and the lower panel controls for a set of sociodemographics and education controls. The findings show that a work capacity limiting health shock increases the probability to receive public WDI in the following year by about 50%. It also increases the probability to not work by almost 50% and reduces individuals' total gross income by €1.7K per year. That coefficient estimates for these six outcomes are statistically significant at the 5% level but we do find evidence that well-being decreases significantly in column (4). The income effect estimate in column (3) represents a relatively modest reduction of less than four percent and does not consider incomes streams by other household members. When considering within-household risk sharing and the equivalized post-tax income, the decrease shrinks to 2.5% and is only marginally significant. It explains why we do not find significant income effects among the notch cohorts in Figure 4 and Table 3 where the identified reduction in DI inflows was much smaller than the identified increase here. The finding also illustrates the ability

[18] We also experiment with interacting the binary D_c variable as in equation (1) with this lagged health shock indicator. However, the limited number of SOEP waves in which the question about health limiting work capacity was asked greatly reduces our statistical power. The interaction term is non-significant at conventional statistical levels in all models and the precision of the main regressor shrinks greatly. Detailed results are available upon request.

of the generous German welfare state to buffer income losses due to health shocks, e.g. through generous paid sick leave policies for short- and long-term sickness (Ziebarth 2013, Burkhauser et al. 2016).

4.4 Reform Impact on Spending and Consumption using EVS data.

Table 4: Impact of 2001 Reform on Quarterly Log Spending and Consumption

	(1) Log (Total Spending)	(2) Log (Private Consumption)	(3) Log (Savings)	(4) Log (Food)	(5) Log (Energy)	(6) Log (Leisure)
D	-0.0816	-0.1361***	-0.4586	-0.1715***	-0.0642	-0.0444
	(0.0708)	(0.0528)	(0.3882)	(0.0566)	(0.0629)	(0.0453)
Gender, age,	X	X	X	X	X	X
State FE	X	X	X	X	X	X
Year FE	X	X	X	X	X	X
Control cohort mean raw	20,543	9,002	1,532	1,065	428	1,189
D	-0.0525	-0.0995**	-0.2106	-0.1061**	-0.0853	-0.0108
	(0.0547)	(0.0488)	(0.3488)	(0.0419)	(0.0629)	(0.0431)
Gender, age	X	X	X	X	X	X
State FE	X	X	X	X	X	X
Year + quarter FE	X	X	X	X	X	X
Socio-dems	X	X	X	X	X	X
Education	X	X	X	X	X	X
Control cohort mean raw	20,543	9,002	1,532	1,065	428	1,189

Notes: SOEP v.33 -- 95% sample. Each column in each panel is from one local cubic polynomial RD model with data-driven bandwidth selection with robust biased corrected inference, see Calonico et al. (2017, 2018, 2019, 2020, 2021). The dependent variables are listed in the column header and are all in logs. All models have 32,188 observations. See main text for more details.

Table 5: Impact of 2001 Reform on Types of Consumption

	(1)	(2)	(3)	(4)	(5)	(6)
	Education	Health	Clothing	Communication	Alcohol & Tobacco	Mobility
D	-0.0169	-0.0891	-0.0165	-0.0101	-0.0036	-0.0697
	(0.0966)	(0.0669)	(0.0426)	(0.0166)	(0.0049)	(0.0669)
gender, age	X	X	X	X	X	X
state FE	X	X	X	X	X	X
Year + quarter FE	X	X	X	X	X	X
Control cohort mean raw	67	239	374	206	892	147
R-squared	0.2080	0.0457	0.0325	0.0056	0.1276	0.0388
D	-0.0213	-0.0788	-0.0164	-0.0096	-0.0038	-0.07229
	(0.0961)	(0.0664)	(0.0423)	(0.0166)	(0.0049)	(0.0792)
gender, age	X	X	X	X	X	X
state FE	X	X	X	X	X	X
year + quarter FE	X	X	X	X	X	X
socio-dems	X	X	X	X	X	X
Education	X	X	X	X	X	X
Control cohort mean raw	67	239	374	206	892	147
R-squared	0.2175	0.0610	0.0475	0.0080	0.1278	0.0400

Notes: SOEP v.33 -- 95% sample. Each column in each panel is from one RD model as in equation (2), enriched with cubic terms. The dependent variables are listed in the column header and are all in logs. All models have 69,543 observations. See main text for more details.

4.5 Demand for Private Insurance Using Private Insurers' Accounting Data

Next, we make use of official data by the *German Association of Insurers (*GDV). These data are an aggregate of the accounting data of all insurers licensed to sell policies in Germany. Specifically, they indicate the inflow of new policies by year and type of insurance. We focus on the market for private ODI, which is one of the biggest markets worldwide.[19] It is almost entirely an individual private market with risk-rated policies and no guaranteed issue. Today, about 26% of all private households hold an individual private DI policy; among households whose household head is an employee the share is 38% (Statistisches Bundesamt 2018).

In addition to the number of new ODI policies sold per year, we also include the number of new *private long-term care* policies, *term life insurance* policies as well as so called "*endowed contract" life insurance* policies. The latter are a popular product in Germany and combine a savings component with a term life insurance policy. In contrast to *term life insurance* policies that only pay out benefits when the policyholders dies, *endowed contract life insurance* policies have a specified contract duration, typically 20-30 years, and pay out a lump-sum either in case of death *or* when the contract terms are up, typically around retirement. As they are not risk-rated, they are a possible old-age savings alternative to private ODI policies.

Figure 5 shows the development of new private insurance policies sold by year and type of policy. The leftmost dotted vertical line indicates the first of a series of public DI reforms, which became effective in 1996 (Burkhauser et al. 2016). The middle dashed vertical line indicates when the 2001 reform was first announced by the center-right government in 1997 (which lost the election in the fall of 1998, leading to a temporary

[19] https://www.gdv.de/de/themen/news/5-fakten-zur-berufsunfaehigkeitsversicherung-34338 (last accessed, March 20, 2021)

suspension of the reform by the new government). The rightmost solid vertical line indicates when the 2001 reform became finally effective after a period of policy uncertainty.[20]

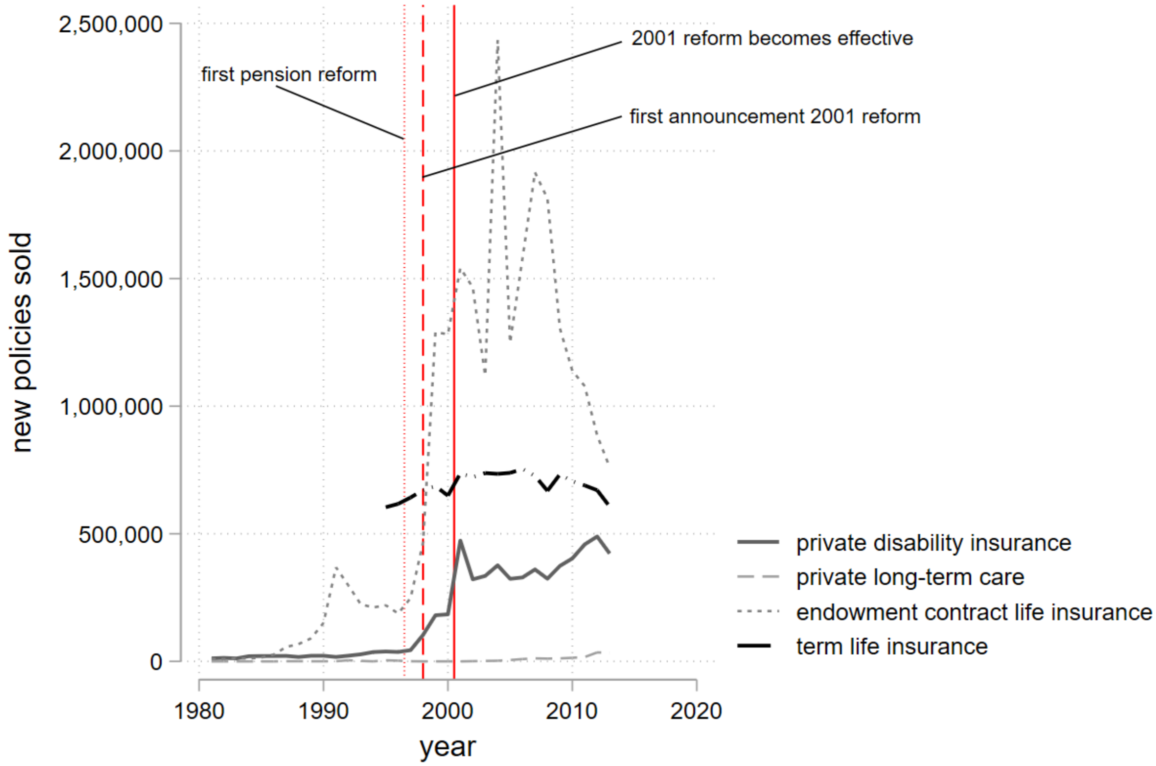

Figure 5: New Private Insurance Policies Sold by Type and Year

As seen, whereas the development of *private long-term care insurance* policies has remained flat at a low level over the entire time period, new *term life insurance* policies were sold at much higher rates of around 700K per year (relative to 82 million German residents). However, their development has also remained relatively flat over the observation period. By contrast, the private ODI market started to take off after the first public DI reform in 1996, and further sharply expanded after the announcement and enactment of the 2001 reform. However, simultaneously, the market for *endowed contract life insurance* policies started to boom, first when

[20] Note that a series of pension reforms increased the statutory retirement age and cut benefits. After a first pension reform in 1992, another passed in 1996 and became effective January 1, 1997. A third major reform package passed in December of 1997 and became effective January 1, 1999 (Rentenreformgesetz 1999). Moreover, follow-up reforms of the new center-left government fundamentally altered the private and public pillars of the pension system; for example, taxation of pensions and endowed contract life insurance policies changed significantly effective January 1, 2005. The spike in 2006 in Figure 2 for endowed life insurance policies is evidence of this.

the first major pension reform was passed in 1992, and even more with the announcement of the 2001 reform (see Figure 1).

Next, we use these data in a motivating simple DD regression framework, similar to equation (1). Here, the dependent variable is the number of new insurance policies sold per year (in 1000s). We define the private ODI market as D_p, our "treatment group." As a first main regressor of interest, we interact D_p with a post 2001 indicator *effective* and, as a second regressor of interest, we interact D_p with a post 1997 indicator *announced*. Each column in Table 4 shows the result of one separate regression with different sets of controls as indicated in the bottom panel.

Table 4: Demand for Private Insurance following Reforms Using Census of Policies

	(1)	(2)	(3)	(4)	(5)
D*effective	201.5850***	201.5850***	201.5850***	78.7346	201.5850***
	(30.2464)	(28.3744)	(11.7643)	(113.9295)	(12.4331)
D*announced	127.9611***	129.8891***	115.0783***	-213.2216	117.7514***
	(25.8238)	(23.5653)	(25.2325)	(262.8053)	(17.2249)
EndowedLife*effective					368.5514***
					(12.4331)
EndowedLife*announced					754.7291***
					(17.2249)
Policy fixed effects	X	X	X	X	X
Year fixed effects		X	X	X	X
Only after 1990			X	X	X
Endowed life insurance				X	X
Pre-treatment mean	22.61094	22.61094	29.93083	29.93083	29.93083
N	85	85	65	88	88
R-squared	0.9877	0.9901	0.9044	0.4117	0.7623

Sources: GDV data 1981-2013, each column stands for one model as in equation (1), estimated by OLS. The outcome variable is new policies sold in 1000s. The columns differ by the sample selection and sets of covariates as indicated in the column headers. *D* stands for the private ODI market and *EndowedLife* for the endowed life insurance market as shown in Figure 5. The lower part of each panel indicates which control variables are included. See main text for more details. Robust standard errors in parentheses are bootstrapped using 1000 replications. *** p<0.01, ** p<0.05, * p<0.1.

The formal regression results confirm the visual evidence in Figure 5. Relative to the market for *private long-term care insurance* and *term life insurance*, the market for private ODI expanded significantly after the announcement of the 2001 reform and, then, after its enactment even more. Specifically, after 2001, about 200K more private ODI policies were sold compared to the other two markets, which is roughly a tenfold(!) increase relative to pre-1997 sales (columns [1]-[3]). However, when adding the market for *endowed life insurance* policies to the regression in column (4), these significant point estimates shrink substantially in size and become insignificant. The reason is simply that the market for life insurance policies with savings component also expanded substantially, even more, over the same time period (column [5]).

4.5 Advantageous Selection and Latent Demand in the Private ODI Market Using SAVE Data

Next, we zoom into the decisions of notch cohorts to purchase private ODI policies, possibly as a substitute for the downgrading of public ODI to WDI. A straightforward hypothesis is that affected private households may have responded to the reform of 2001 by an increased demand of private ODI insurance—in other words, the reform may have crowded-in demand for private ODI. A rich literature has studied the reverse effect, the crowd-*out* effects of private health insurance by public health insurance expansions (Section 2); this paper is one of the first to test for this link in the market for DI.

To do so, we rely on representative survey data from the SAVE survey (*Sparen und AltersVorsorgE in Deutschland*, “Saving for Old Age in Germany”*),* Coppola and Lamla (2013) provide a detailed overview of the dataset. The SAVE data are unique because they include a very rich set of questions regarding risk aversion, expectations about individual life expectancy, health, and risk aversion. These are precisely the predictors of insurance demand that are typically unobserved by researchers and insurers, resulting in asymmetric information. In our case, these unique survey measures help us to (a) mimic the risk classification system of private DI insurers to identify households for whom private policies are unaffordable, or whom insurers entirely deny coverage. Moreover, they help us to (b) measure private information that drive insurance market selection in the spirit of Akerlof (1970) and Hendren (2017). Measures of subjective life expectancy, risk aversion, savings attitudes and disposable household income help us to characterize latent household demand for insurance, which is crucial to assess potential welfare effects of public DI reforms, but also potential regulation of the private market such as guaranteed issue or community rating.

We use all existing waves of the SAVE dataset, which was conducted biannually, from 2001 to 2010. Moreover, we ignore respondents below the age of 20 and those above the age of 64. Table A4 shows the summary statistics of our main sample. The main variable of interest is the binary *Private ODI,* which indicates whether the households holds a private ODI policy. This question was asked in all waves and 36% of all households are ODI policyholders. Moreover, to identify the notch cohorts, it is crucial to know the birth year, which we directly observe as a separate variable. We run the same RD model as in equation (2) above, but focus on linear specifications as they fit the data better.[21]

Figure 6 below shows the main visual result for four different specifications: the full sample, solely respondents who are eligible for a public pension (civil servants and the self-employed are not), solely non-married respondents, and solely single households (clockwise starting from the left upper subgraph). Following Figure 3 and 4 above, the x-axes display the birth year and the y-axes display the outcome y_{it}.

We find the following: First, maybe surprisingly, the demand slope is clearly and strongly increasing in the birth cohorts. In other words, younger people are much more likely to hold a private ODI policy in Germany. This observation is not surprising on second thought, however, and entirely in line with Figure 5 above. The reason is that, after a strong expansion of the welfare state in the decades after WWII (especially in the 1970s), German policymakers started to implement a series of structural reforms of the statutory pension and DI system in the 1980s and particular 1990s. These were accompanied will strong messaging, education (even in high schools) and lobbying that private insurance policies for old age protection were crucial for young people. In addition, younger cohorts are much less likely to be rejected by private ODI insurers and offered lower premiums as they are healthier and have fewer pre-existing conditions.

Second, in the market equilibrium, none of the four graphs in Figure 6 shows an obvious discontinuous jump in the likelihood to have private ODI insurance for the notch cohorts. While single insurers may certainly have specifically targeted subgroups that were affected by the 2001 reform, representative data do not yield any evidence for a systematic crowding-in or substitution effect. This finding is consistent with the census of market-level data in Figure 5 as it is extremely unlikely that the entire tenfold increase in market demand is driven by the notch cohorts alone, and as we see strong demand increases already after the first 1996 DI reform which did

[21] The cubic variants are available upon request.

not target the notch cohorts investigated in this paper. It is also in line with the literature on incomplete financial knowledge (Lusardi et al. 2017) as well as German surveys according to which a majority of Germans, especially younger people, is not aware of the specifics of the 2001 reform (Metallrente 2020).

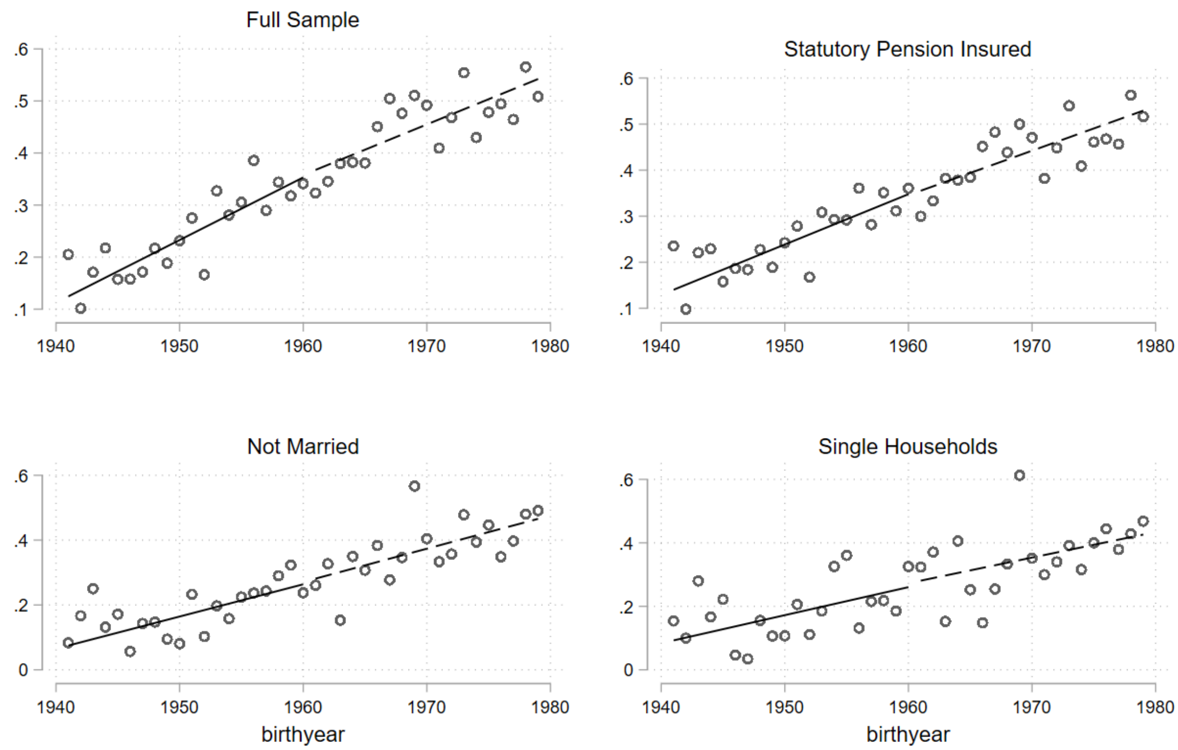

Figure 6: Private Disability Insurance by Birthyear

Table 5 shows the equivalent parametric RD results for what we see in Figure 6, estimating again equation (2). The upper panel shows the results for the four samples in Figure 6 when only controlling for the clearly exogenous age, gender, state and year fixed effects. The lower panel shows the results when adding rich sets of controls for socio-demographics, education and the labor market. As seen, all eight models confirm the visual evidence in Figure 6: all eight point estimates are small relative to the pre-treatment mean, the signs of the estimates are not consistent and alternate between being positive and negative. None of the eight RD estimates is statistically different from zero; they vary between -0.008 and 0.013 and let us conclude that we do not find evidence for our crowding-in hypothesis when it comes to the notch cohorts specifically. The strongly increasing age-gradient should be kept in mind, however.

Table 5: Affected 2001 Notch Generation and Demand for Private Disability Insurance

	(1) Full	(2) SPI insured	(3) No married	(4) Single HH
D_i	0.011	0.007	0.013	0.003
	(0.0236)	(0.0240)	(0.0389)	(0.0477)
$(1-D_i)$*yob_i	0.012***	0.011***	0.010***	0.009***
	(0.0011)	(0.0014)	(0.0023)	(0.0027)
D_i*yob	0.007***	0.007***	0.007***	0.006*
	(0.0018)	(0.0019)	(0.0024)	(0.0028)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
Pre-treatment mean	0.2653	0.2689	0.1949	0.1905
N	12,015	9,566	4,149	2,136
R-squared	0.065	0.056	0.069	0.074
D_i	0.006	0.001	-0.008	-0.010
	(0.0181)	(0.0203)	(0.0294)	(0.0454)
$(1-D_i)$*yob_i	0.011***	0.010***	0.007***	0.006**
	(0.0013)	(0.0016)	(0.0019)	(0.0029)
D_i*yob	0.010***	0.009***	0.007***	0.006**
D_i	(0.0013)	(0.0014)	(0.0020)	(0.0029)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
socio-dems	X	X	X	X
education	X	X	X	X
employment	X	X	X	X
Pre-treatment mean	0.2653	0.2689	0.1949	0.1905
N	12,015	9,566	4,149	2,136
R-squared	0.114	0.107	0.138	0.146

Sources: SAVE data 2001-2010, each column in each panel stands for one model as in equation (1), estimated by OL. The columns differ by the sample selection as indicated in the column headers. The lower part of each panel indicates with control variables are included. See main text for more details. Robust standard errors in parentheses are clustered at the running variable, distance of year-of-birth (yob) to the eligibility cut-of 1961. *** p<0.01, ** p<0.05, * p<0.1.

Our interpretation of the overall evidence is that the entire series of structural pension and DI reforms—combined with the messaging of policymakers and the private insurance industry—substantially increased the general demand for private ODI at the time, but also for private life and pension insurance with savings component. But there is no evidence that the notch cohorts specifically responded.

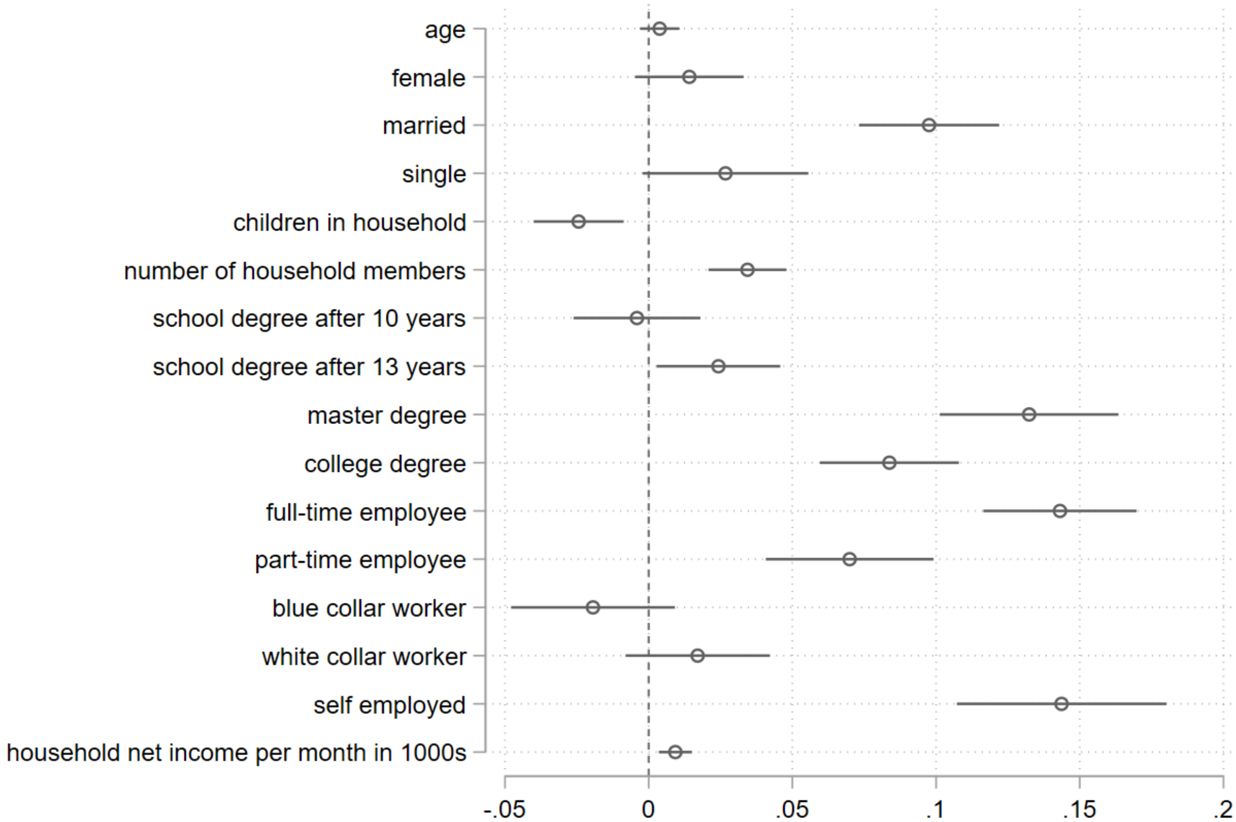

Figure 7: Socio-Demographic Demand Determinants of Private ODI

In a next step, we investigate the demand for private ODI in more detail. We do so in order to understand the market better and be able to assess potential welfare effects of structural DI reforms. Figure 7 plots the coefficients of a multivariate regression of *Private ODI* on X'_{it} where Table A4 lists the means of the socio-demographics included. As seen, when controlling for all other background variables, age and gender are not significant demand predictors, but being married and being a larger household predicts the likelihood to hold a private ODI policy significantly. Moreover, the self-employed buy private ODI policies are significantly higher rates, as do higher income households. In addition, being a full-time employee and higher education increase the likelihood of holding private ODI policies significantly—these effect sizes are large. All these relevant demand predictors are entirely in line with expectations and economic intuition.

Next, Figure 8 takes advantage of a series of health status, health behavior and health care utilization measures (Table A4). Measuring health is crucial in this setting as private ODI policies are individually underwritten and unhealthy applicants may not obtain affordable policies or are denied coverage by insurers. Hence our expectation is that sick individuals are less likely to hold private ODI policies. These expectations are confirmed by Figure 8, which plots the relevant health coefficient but simultaneously controls for age, age squared and gender. As seen, respondents with a high health satisfaction are significantly more likely to hold private ODI policies. By contrast, those who assess their own future health very negatively, who are current smokers and spent many nights in a hospital in the past 12 months are significantly less likely to hold a policy.

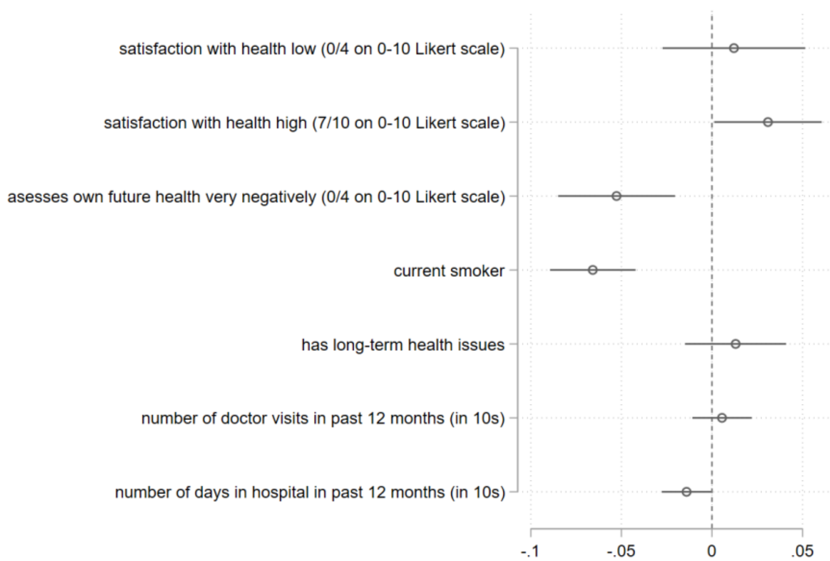

Figure 8: Advantageous Selection in the Private ODI Market

In a next step, we use a principal component analysis to summarize and aggregate all available objective and subjective health measures into a continuous health risk index (Jolliffe 2002). The distribution of this standardized risk index with mean 0 is in Figure A2. It is reassuring to see and obtain a typical left-skewed health risk distribution with a long right rail (cf. Karlsson et al. 2016).

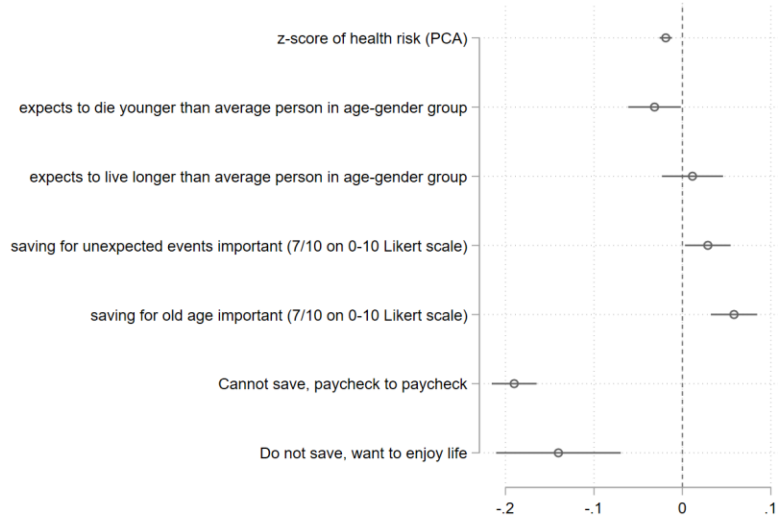

Figure 9: Private Information Predictors of Private ODI Demand

In the next step, we regress y_{it} again on age and gender, this continuous health risk measure and other, typically unobserved demand predictors. Figure 9 clearly shows that an increase in the z-health risk score by one unit has a highly significant and tight negative predictive impact on the probability to hold a private ODI policy. (Note that the scale of the health risk score differs from the other predictors and ranges from -2.1 to +14.1, Table A4) This confirms our prior and is in line with the notion that unhealthy applicants are denied coverage (or do not even apply for) private ODI.

This finding is also in line with a survey of new public WDI recipients from 2008 (Märtin et al. 2012; 2014). While the overall share of respondents with private ODI exceeds 25% and is even higher among employees, the share of WDI recipients who held private ODI is a mere 6% (among those, 87% actually received benefits). Regarding reasons for not owning a private ODI policy, about 29% of respondents say they could not purchase a policy due to pre-existing conditions, and 55% claimed a policy was too expensive.[22]

The evidence from the other, typically unobserved, private information measures also reinforce common priors and illustrate why policyholders in the Germany private ODI market are a select sample: As seen in Figure 9, controlling for health, respondents who expect to die younger than the average person in their age-gender

[22] The share of people without or with unknown educational degree is 35%, another 33% has a vocational degree. Only about 3% have a university degree. This is another indicator for the negative selection of this group.

group are significantly less likely to purchase policies. This finding mirrors rational demand, as the expected value of disability pensions increases in the life expectancy, conditional on health. Figure A3 further decomposes the reasons for this subjective expectation and shows that knowledge about future diseases or an unhealthy lifestyle are dampening the demand and result in an advantageously selected market, in line with Soika (2018). By contrast, when individuals believe that they will live significantly longer than the average person with their age and gender, they are significantly *more likely* to hold a private ODI policy.

Figure 9 also very clearly shows that saving attitudes and liquidity constraints matter. Respondents who find it very important to save for unexpected life events and old age are much more likely to purchase private ODIs. By contrast, liquidity constrained households who say that they live from paycheck to paycheck are *substantially* less likely to be insured, as are those who prefer to live in the moment and have fun. The effect sizes of the last two predictors are very large.

In a very last step, we try to characterize and cluster uninsured households. As implied by the previous discussion, there could be several reasons for why households have no private ODI policy. One is the inability to obtain coverage due to risk ratings and the absence of guaranteed issue; another is private knowledge about health factors and life expectancy, and yet another are risk preferences and low demand because the public safety net option may be sufficient for some. Naturally, liquidity constraint households may not be in a position to purchase the relatively expensive policies, which cost on average XX in 2018. When applying optimal k-means clustering on a number of relevant demand predictors from Figures 7 to 9, we find that a cluster of three groups is optimal (see Figure A4 and Makles, 2012). These three clusters allow us to characterize the latent demand among those households who did *not* purchase a private ODI.

Group 1 consists of about a quarter of the two thirds of households without coverage. They are relatively young with average age 34 but also relatively sick and low-income, which explains why they do not have private policies. Their health risk score is below average and negative, and their income the lowest of all three groups.

Group 2 consists of a third of those without private ODI coverage. This is by far the group with the oldest age, on average 58 years. Although—or precisely because—this group is relatively healthy (the healthiest on average), the old age and their proximity to retirement (early retirement starts for some population subgroups from age 60) is likely one main reason for their low private ODI demand. This prior is reinforced as this group has the highest share of respondents who say that, for them, "saving for old age is relatively unimportant."

Finally, Group 3 is the biggest group and makes up almost half of all uninsured households. They are between 41 and 52 years old and relatively healthy, although half of them have negative health risk scores. When also flagging those 29% in that group who say that they have no possibility to save any money because they live paycheck to paycheck, we end up with a relatively small group of relatively healthy middle-aged individuals who have relatively high household incomes. The have latent private ODI demand and would be a fruitful target group for private insurers and policymakers who intend to increase the share of households with private coverage.

5. Conclusion

This paper comprehensively studies the effect of a fundamental reform of the German Public Disability Insurance System using a range of outcome margins, administrative and survey data. The reform became effective in 2001 and changed the public DI eligibility requirements for "notch cohorts" born after 1960. These cohorts were no longer eligible to receive a public disability pension when their health status prevented them from working in their previous occupation. Instead, they only become eligible for full disability pensions if their health prevents them to work for at least three hours in *any* occupation.

Using administrative data and the universe of public DI inflows, we first show that the reform significantly reduced the inflow of recipients by about 20% for men and 10% for women. We then confirm these findings by using a representative sample of the underlying population of interest. Using SOEP household data, we confirm that the notch cohorts are significantly less likely to receive public DI pensions. At least at the population level, when studying labor supply, we do not find statistically significant increases in the probably to either leave the labor force or work full-time. However, we do find some evidence for lower subjective well-being among affected notch cohorts.

Next, we study the impact of the decrease in public DI generosity on the demand for private Occupational Disability Insurance (ODI). Germany has one of the biggest private ODI markets in the world. However, these are individually underwritten policies, which are health risk-rated—guaranteed issue does not exist in this market. We provide evidence that a series of structural reforms at the time increased the demand for private old age insurance substantially. However, using again representative data, we do not find that the demand for private ODI increased significantly among the notch cohorts. Next, we carve out several ODI demand predictors, some of which are typically unobserved and contribute to asymmetric information. Such unobservables are private beliefs about individual life expectancy being higher or lower than in actuarial life tables. We can even pinpoint whether respondents believe that they will die younger due to worse genes, an unhealthy lifestyle or their health. We also observe measures of risk aversion, attitudes toward savings and whether the household is liquidity constrained.

We find that better educated and larger households are more likely to hold a private ODI policy, as are married people and those who are self-employed. By contrast, smokers and those who view their future health very negatively are substantially less likely to be covered. Using a series of health measures, we create a standardized health risk measure, which is a highly significant predictor of not having private coverage. Expecting to die young due to a disease or a risky lifestyle makes it significantly less likely to hold a private ODI policy as well; the same is true for liquidity constrained households and those who say that they prefer to have fun now rather than to care about the future. Conversely, believing that it is very important to save for old age and unexpected events is highly predictive of being insured. In a final step, we categorize and characterize uninsured households into three groups: (i) the old with low demand due to being close to retirement, (ii) the young and sick whose risks exceed their willingness to pay, and (iii) a group of middle-aged whose willingness to pay is close to their risk.

The final part of the paper uses rich consumer expenditure data to estimate the impact of the 2001 reform on household consumption (TO BE WRITTEN).

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Appendix

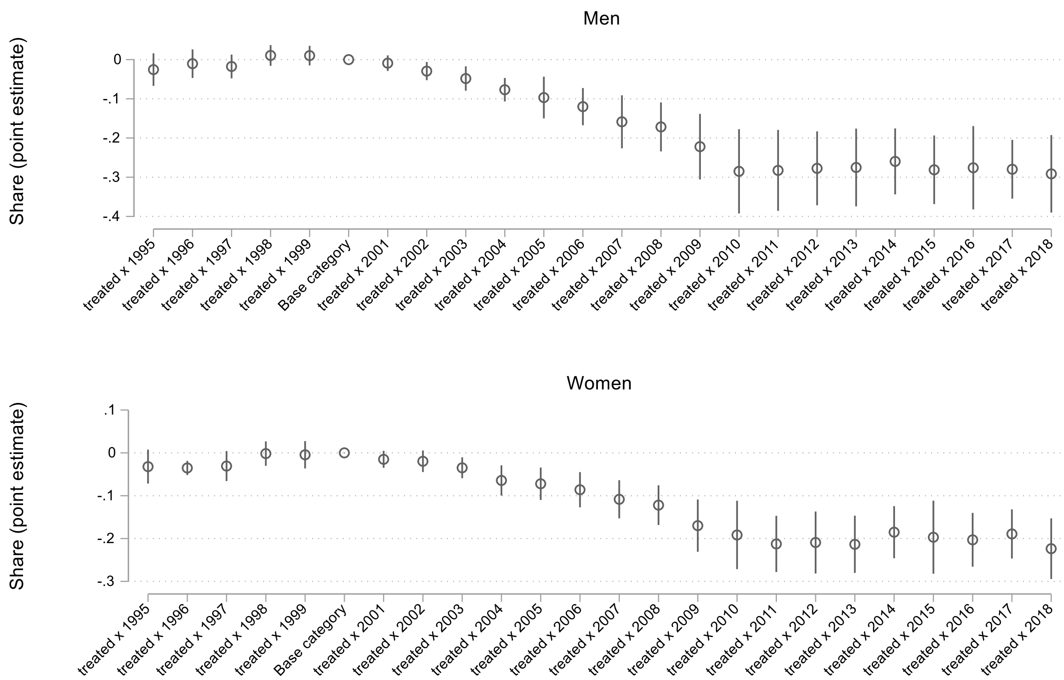

Figure A1: Effect of 2001 Reform on Public DI Inflows—Event Study Men and Women

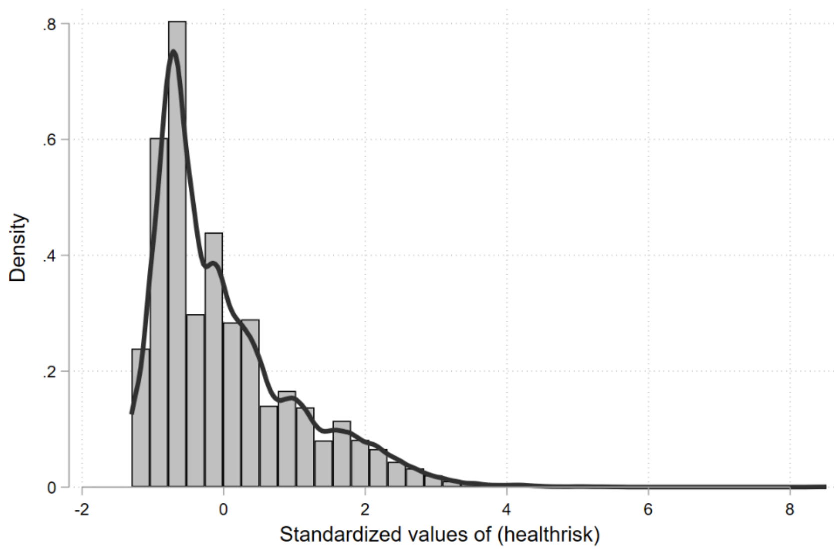

Figure A2: Estimated Standardized Risk Score

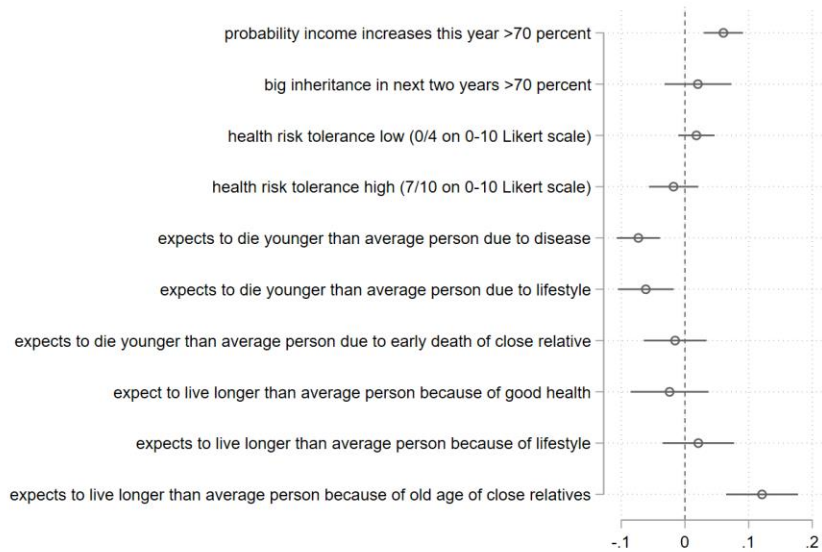

Figure A3: Other Demand Predictors of Private ODI Policies

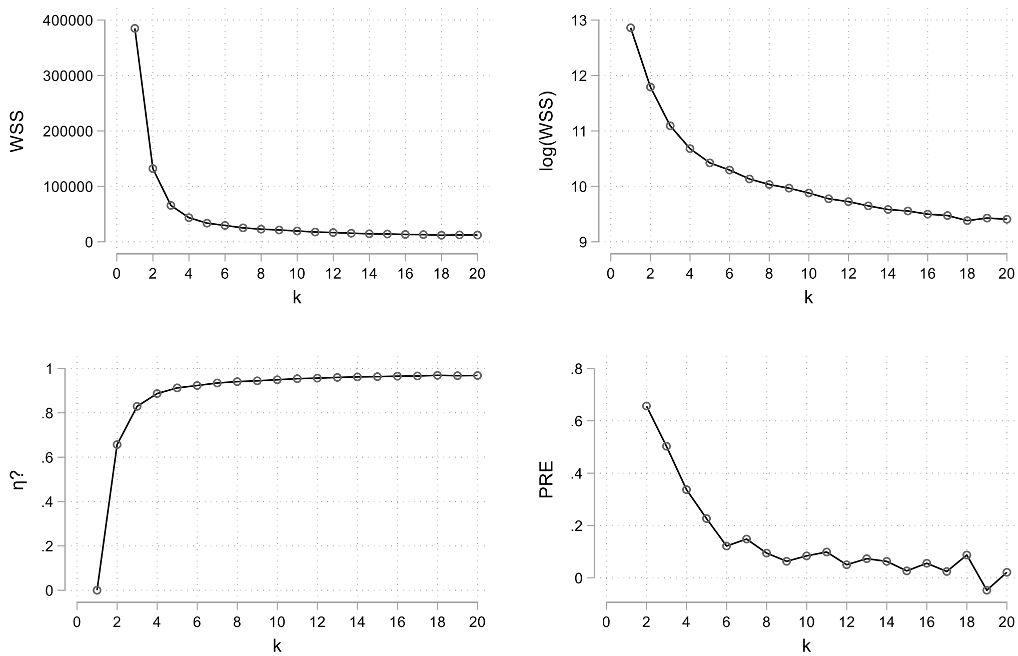

Figure A4: Key Statistics for Optimal K Cluster Solutions of Demand Predictors

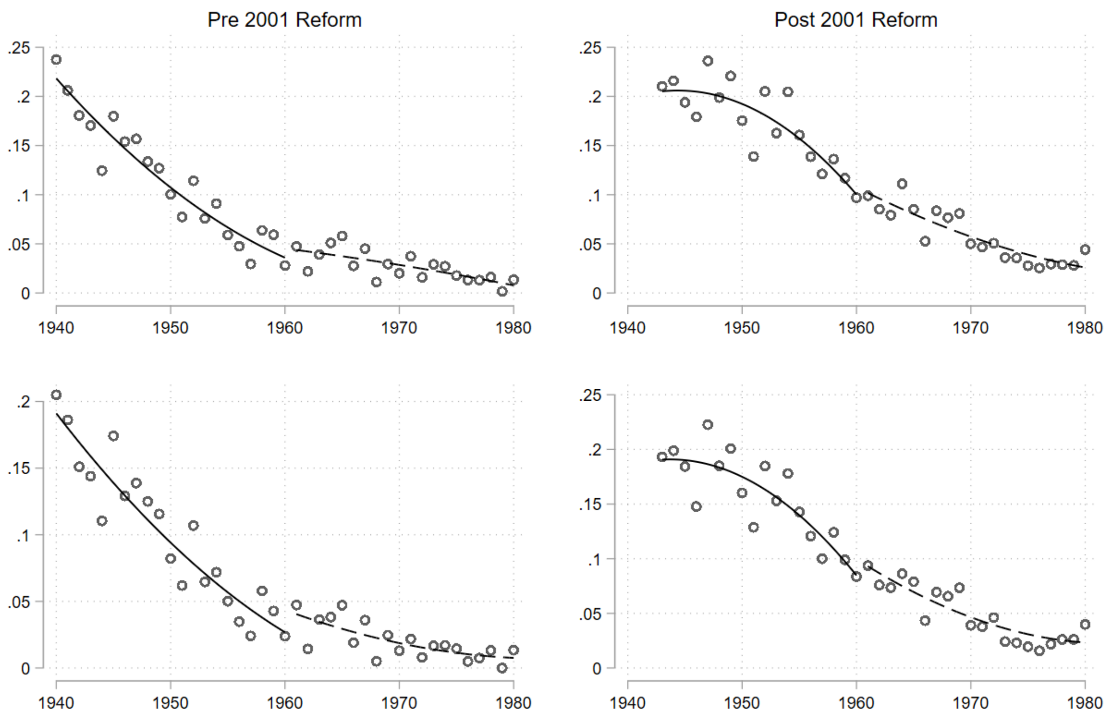

Figure A5: Being Handicapped (left) and Officially Disabled (right) by Birthyear

Table A1: Descriptive Statistic, SOEP Data, 1995-2016

		Mean	SD	Min	Max	N
	Public DI I	0.0344	0.1822	0	1	119784
	Public DI II	0.0289	0.1676	0	1	119784
	Handicapped	0.0867	0.2814	0	1	119784
	Officially disabled	0.0752	0.2637	0	1	119784
	Non employed	0.1401	0.3471	0	1	119784
	Full-time employed	0.8025	0.3981	0	1	119784
	Individual total income *(equivalized)*	35.7573	33.6449	0	2580	119784
	Subjective well-being	6.9709	1.7503	0	10	119784
Socio-demographics						
	Age	42.2895	10.0663	17	59	119784
	Age squared	1890	830	289	3481	119784
	Married	0.6488	0.4774	0	1	119784
	Single	0.2419	0.4283	0	1	119784
	Children in household	0.8571	1.0696	0	11	119784
	Adults in household	0.3697	0.6948	1	7	119784
	Household size	2.2268	1.1855	1	13	119784
	Dropout	0.0258	0.1585	0	1	119784
	Schooling degree 9 yrs	0.2838	0.4508	0	1	119784
	Schooling degree 10 yrs	0.2973	0.4571	0	1	119784
	Schooling degree 13 yrs	0.2196	0.4140	0	1	119784
	Civil servant	0.0677	0.2513	0	1	119784
	Self-employed	0.1097	0.3125	0	1	119784
	White collar	0.3473	0.4761	0	1	119784
	Public Sector	0.1693	0.3750	0	1	119784
	Part-time employed	0.0415	0.1996	0	1	119784
	In job training	0.0144	0.1191	0	1	119784
Health						
	Health Satisfaction	6.8714	2.0976	0	10	119155
	SAH	2.4687	0.8974	1	5	119633
	Publicly insured	0.8235	0.3813	0	1	119784
	Privately insured	0.1633	0.3697	0	1	119784
	# nights hospital	1.0298	6.3872	0	360	119784
	# doctor visits	1.8697	3.6866	0	99	117583

Notes: SOEP v.33 -- 95% sample.

Table A2: Impact of 2001 Reform on Public DI Recipiency Rates: 1995-2000

	(1) Public DI I	(2) Public DI II	(3) Non-Married	(4) Single Household
D	-0.008	-0.010	0.019*	-0.026
	(0.0068)	(0.0064)	(0.0108)	(0.0168)
(1-D)*yob	0.022***	0.015***	0.013***	0.033***
	(0.0032)	(0.0024)	(0.0033)	(0.0069)
D*yob	0.016***	0.010***	0.010***	0.023***
	(0.0028)	(0.0020)	(0.0017)	(0.0046)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
R-squared	0.044	0.034	0.056	0.049
D	-0.005	-0.008	0.013	-0.012
	(0.0059)	(0.0057)	(0.0164)	(0.0156)
(1-D)*yob	0.009***	0.005**	0.009**	0.018***
	(0.0024)	(0.0022)	(0.0037)	(0.0053)
D*yob	0.006***	0.002	0.008**	0.013***
	(0.0021)	(0.0019)	(0.0035)	(0.0040)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
socio-dems	X	X	X	X
education	X	X	X	X
R-squared	0.194	0.162	0.174	0.235
Observations	87,472	87,472	29,109	31,889

Notes: SOEP v.33 -- 95% sample. Each column in each panel is from one RD model as in equation (2), enriched with cubic terms. The first two columns use the full sample, whereas the third column selects on non-married individuals and the final column on single households. The dependent variable is Public DI I, except in the second column where it is Public DI II. See main text for more details.

Table A3: Work Limiting Heath Shocks, Labor Market Outcomes and Subjective Well-Being

	(1) Public DI	(2) Non-Employed	(3) Individual Total Income	(4) Subjective Well-Being
Work Limiting Health Shock	0.0248**	0.0356**	-1.7146**	0.0061
(last year)	(0.0122)	(0.0177)	(0.6807)	(0.0957)
age	-0.0096**	-0.0076	5.4107	-1.5304***
	(0.0041)	(0.0062)	(5.6017)	(0.3316)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
individual FE	X	X	X	X
R-squared	0.0076	0.0116	0.0191	0.0046
Work Limiting Health Shock	0.0246**	0.0353**	-1.7265**	0.0049
(last year)	(0.0122)	(0.0176)	(0.6788)	(0.0957)
age	-0.0068*	0.0066	4.9423	-1.5078***
	(0.0035)	(0.0189)	(5.1619)	(0.3595)
gender, age	X	X	X	X
state FE	X	X	X	X
year FE	X	X	X	X
individual FE	X	X	X	X
socio-dems	X	X	X	X
education	X	X	X	X
Pre-treatment mean	0.04915	0.1057	44.75	7.14
N	14,668	14,668	14,668	14,668
R-squared	0.0082	0.0144	0.0196	0.0055

Notes: SOEP v.33 -- 95% sample. Each column in each panel is from one multivariate regression with the dependent variable listed in the column header. The main regressor of interest is *Work Limiting Health Shock* which is lagged and was elicited in 2011-2013 and 2015 only. See main text for more details.

Table 4: Impact of 2001 Reform on Household Income, Spending and Consumption

	(1) Gross HH Income	(2) Net HH Income	(3) Total HH Spending	(4) Private HH Consumption	(5) HH Savings	(6) Loan Repayment
D	-83.01	-78.66	-323.56*	-49.79	-78.74	-80.16
	(96.96)	(59.97)	(190.94)	(49.19)	(64.81)	(53.84)
gender, age	X	X	X	X	X	X
state FE	X	X	X	X	X	X
Year + quarter FE	X	X	X	X	X	X
Control cohort mean	8113	5971	9810	4332	806	626
R-squared	0.024	0.2937	0.0420	0.0744	0.0082	0.0096
D	-115.64	-78.66	-367.61**	-57.89	-87.83	-86.08
	(84.31)	(59.97)	(180.52)	(46.81)	(64.08)	(53.49)
gender, age	X	X	X	X	X	X
state FE	X	X	X	X	X	X
year + quarter FE	X	X	X	X	X	X
socio-dems	X	X	X	X	X	X
education	X	X	X	X	X	X
Control cohort mean	8113	5971	9810	4332	806	626
R-squared	0.3509	0.2937	0.1330	0.1523	0.0227	0.024

Notes: SOEP v.33 -- 95% sample. Each column in each panel is from one RD model as in equation (2), enriched with cubic terms. The dependent variables are listed in the column header. All models have 69,543 observations. See main text for more details.

Table A4: Descriptive Statistic SAVE Data, 2000-2010

	Mean	SD	Min	Max	N
Private ODI	0.3601	0.4801	0	1	12015
Socio-demographics					
Age	43.9452	10.0271	20	64	12015
Female	0.5107	0.4999	0	1	12015
Married	0.6547	0.4755	0	1	12015
Single	0.2007	0.4005	0	1	12015
Children in household	0.9064	1.0533	0	8	12015
Household size	2.6851	1.2781	1	13	12015
Schooling degree 10 yrs	0.2733	0.4457	0	1	12015
Schooling degree 13 yrs	0.4349	0.4958	0	1	12015
Master degree	0.2619	0.4397	0	1	12015
College degree	0.6163	0.4863	0	1	12015
Full-time	0.5614	0.4962	0	1	12015
Part-time	0.1434	0.3505	0	1	12015
Blue collar	0.2037	0.4028	0	1	12015
White collar	0.3877	0.4872	0	1	12015
Self employed	0.0817	0.2740	0	1	12015
Household net income (in 000s)	2.5470	2.5094	0	120	12015
Health					
Health satisfaction 0-4/10	0.1872	0.3901	0	1	6409
Health satisfaction 8-10/10	0.5907	0.4917	0	1	6409
Concerns about own health	0.2035	0.4026	0	1	6409
Smoker	0.3313	0.4707	0	1	6409
SAH	2.3447	0.7892	1	5	6409
Serious Health Issues	0.4133	0.4925	0	1	6409
# doctor visits	0.5298	0.7295	0	9	6409
# days hospital	0.1403	0.6822	0	27	6409
Health risk score	0.0360	1.6694	-2.15062	14.07128	6409
Expectations and attitudes					
Subjective life expectancy low	0.1830	0.3867	0	1	6409
Subjective life expectancy high	0.1231	0.3286	0	1	6409
Savings 4 Unexpected Important	0.7193	0.4494	0	1	6409
Savings 4 OldAge Important	0.7514	0.4322	0	1	6409
No savings possible	0.2091	0.4067	0	1	6409
No savings, enjoy life	0.0204	0.1415	0	1	6409
Higher income expected	0.1317	0.3382	0	1	7815
Inheritance expected	0.0384	0.1921	0	1	7815
Health private information					
Health risk tolerance 0-4/10	0.7301	0.4439	0	1	7815

Health risk tolerance 8-10/10	0.1179	0.3225	0	1	7815
Expects disease shortens life	0.0984	0.2979	0	1	7815
Expects lifestyle shortens life	0.0600	0.2375	0	1	7815
Expects death of relative shortens life	0.0470	0.2116	0	1	7815
Expects good health increases life	0.0495	0.2170	0	1	7815
Expects lifestyle increases life	0.0612	0.2396	0	1	7815
Expects relatives genes increases life	0.0458	0.2091	0	1	7815

Sources: SAVE data 2001-2010.