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Abstract:

Disability insurance (DI) systems are widely criticized for their inherent work disincentives. This paper evaluates the effects of a Swiss DI reform that aims to lower pensions for a group of existing DI beneficiaries and introduces an additional notch to the pension schedule. The reform does not significantly affect average earnings and employment, but increases the disability degree of those threatened by a pension decline. We estimate bounds on the income and substitution effects employing the principal stratification framework. The in-come effect is quantitatively important, while the substitution effect is smaller and bounds include zero.

Keywords:

Disability insurance, work disincentives, income and substitution effects, partial benefit system.

JEL Classification:

C30, I13, J01.

1 Introduction

High numbers of people with disabilities, their low labor market attachment and high dependency on social assistance create considerable costs to society (OECD, 2010). Many countries are thus forced to reform their disability insurance (DI) systems. One of the prime problems that needs to be solved are work disincentive effects of DI systems: Because DI pensioners fear losing a significant part of their benefits if labor supply exceeds certain thresholds – so called "cash cliffs" – they do not raise employment above this level (substitution effect). Furthermore, DI benefits increase non-earned incomes, which reduces employment if people prefer leisure over labor (income effect). While these two mechanisms are well understood in theory, identifying income and substitution effects empirically is challenging. Individual reactions to changes in the benefit schedule usually reflect both mechanisms jointly and are therefore not informative on either effect.

This paper presents novel insights on the importance of the income and substitution effects by evaluating a reform of the Swiss DI system and employing a principal stratification approach. The Swiss system insures partial disability, where beneficiaries can work and claim DI benefits at the same time. The amount of DI benefits is a step-wise function of the disability degree, which is assessed by the disability insurance and denotes the presumed earnings loss due to the disability (in percent). In January 2004, Switzerland further graduated the pension system and introduced a three-quarter pension, additionally to the already existing quarter pension, semi pension, and full pension. The reform led to a substantial loss in DI benefits for a subset of beneficiaries and imposed a new earnings threshold for a full pension. The income effect increases labor supply because the loss in pension needs to be compensated by an increase in earnings. The substitution effect reduces the incentives to work, because a reduction in earnings signals an increase in the disability degree and therefore can lead to a preservation of the full pension.

We first evaluate the total effect of the reform on employment and earnings using a local difference-in-differences approach taking advantage of the sharp discontinuity that separates individuals who are fully exposed to the reform (born after December 31, 1953) and individuals who were exempted from the benefit cut (born before December 31, 1953). We then use the principal stratification framework (Frangakis and Rubin, 2002) to decompose the total effect, and provide bounds for income and substitution effects. The resulting bounds can be sequentially tightened by adding revealed preference restrictions motivated by a simple static labor supply model.

We find a small total effect on employment (1 to 2 percentage points), and no effect on earnings. Decomposing the total effect, we find informative bounds for the income effect. The reform increased employment for individuals actually losing 25% of their DI benefits by 9 to 20 percentage points and earnings by 136 to 3135 CHF, which is up to 50% of mean pre-reform earnings. Bounds on the substitution effect are smaller and suggest only a small labor supply reaction to the reform of individuals who kept a full pension. Even though the substitution effect is modest, we still find an immediate and persistent increase in the disability degree of approximately 3 percentage points. This is driven by the fact that around 75% of the targeted beneficiaries managed to increase their disability degree and thus keep their full pension, many of them without even being forced to reduce labor supply.

There is a relatively large literature on work disincentives imposed by the disability insurance, which has in common that implicit or explicit changes of the budget constraint are used to derive structural parameters of labor supply (see Bound and Burkhauser, 1999 for a review). The existing literature exploits reforms on the generosity of the DI system (Campolieti and Riddell, 2012; Gruber, 2000; Kauer, 2014; Kostol and Mogstad, 2014; Marie and Vall Castello, 2012; Schimmel et al., 2011; Weathers and Hemmeter, 2011) or on eligibility criteria (Autor and Duggan, 2007; Borghans et al., 2012; Karlström et al., 2008; Moore, 2014; Staubli, 2011), or compares labor supply of accepted and rejected DI applicants (Bound, 1989; Chen and van der Klaauw, 2008; French and Song, 2014; Maestas et al., 2013; von Wachter et al., 2011). Overall, the finding suggest that the DI system imposes work disincentives to some individuals, in addition to providing income to individuals who are at need.

Our study differs in three important ways from the previous literature: First, to the best of our knowledge there is no literature on the performance of a partial DI benefit system, even though many countries (such as France, Germany, Netherlands, Spain, Sweden and Switzerland) already rely on partial DI systems and there is also increasing interest in the US to provide partial income support (Autor and Duggan, 2010). Second, with the exception of two studies in non-work contingent systems (Autor and Duggan, 2007; Marie and Vall Castello, 2012), the previous literature cannot distinguish between income and substitution effects, and typically predicts total effects only. A better understanding of the importance of these two effects is crucial for designing effective DI systems. Third, we develop a novel empirical framework motivated by a labor market model that allows deriving bounds on these conflicting effects. This methodology relates to Kline and Tartari (2016), who study bounds of labor supply responses to the US job first program within a revealed preferences framework.

The paper proceeds as follows. Section 2 provides details on the Swiss DI Act. Section 3 outlines the expected effects of the reform. Section 4 discusses the empirical identification strategy. Section 5 presents the data and descriptive statistics, section 6 shows the results, and section 7 concludes.

2 Swiss Disability Insurance Act

In Switzerland, the mandatory public DI insures individuals against the partial or full loss of the ability to work due to impaired health. DI is permanent, i.e., an individual can claim disability benefits for as long as the health condition is unchanged. Overall, the system is very generous. The public DI system together with the occupational pension scheme guarantee replacement rates of at least 60% of previous earnings, in most cases much higher.

The Swiss DI system allows for partial disability. DI benefits are a stepwise function of the disability degree. The disability degree denotes the presumed earnings loss due to the disability (in percent) and is determined by the DI office. In particular, the disability degree is assessed by the caseworkers in the following manner:

$$\text{disability degree (dd)} = 1 - \frac{\text{potential earnings with disability}}{\text{potential earnings without disability}}$$

Typically, *potential earnings without disability* are predicted on the basis of the individual's earnings before disability onset, and *potential earnings with disability* on the basis of the individual's earnings during disability. This procedure is only valid if the DI beneficiary exhausts his or her remaining work capacity. If the caseworker concludes that the person has idle work capacity – for example because medical records suggest that the person could work a higher number of hours, he can fix potential earnings based on assumed work capacity and official wage indices.

The fourth revision of the Swiss DI Act introduced the three-quarter pension for individuals with a disability degree between 60 and 70%, additionally to the already existing quarter-, semi-, and full pensions (see table 1). This implies for individuals with a disability degree between 60 and 66% that they gain a quarter of their DI pension, and for individuals with a disability degree between 67 and 69% that they loose a quarter of their DI pension if their disability degree remains unchanged. Furthermore, it introduces two new disability degree notches at 60% for a three-quarter pension and at 70% for a full pension. In this paper we evaluate behavioral responses to a potential reduction in DI spendings by focussing on individuals with an initial disability degree between 67 and 69%. For this subgroup the reform

Table 1: The partial DI system (before/after the reform)

Pension	before Jan 1, 2004	since Jan 1, 2004
Full	$67\% \leq$ dd	$70\% \leq$ dd
Three-quarter	none	$60\% \leq$ dd $< 70\%$
Semi	$50\% \leq$ dd $< 67\%$	$50\% \leq$ dd $< 60\%$
Quarter	$40\% \leq$ dd $< 50\%$	no change
No pension	dd $< 40\%$	no change

Note: dd stands for disability degree and is potential earnings loss due to disability in percentage to earnings potential without a disability.

provides a suitable control group as individuals who were older than 50 years in January 2004 were exempted from the pension cut.[1]

The fourth revision of the Swiss DI Act was already planned in the late 1990s. Lobbyists were able to request a referendum and the Swiss people disapproved of the reform in 1999 with an unusual high no-share of 70%. There is a wide consensus that the failure of the original reform was mainly due to the planned abolition of the quarter pension level. The federal government adjusted the law accordingly and kept the quarter pension. The introduction of the three-quarter pension was first discussed in 2002 and passed the parliament in 2003. Since no referendum was requested, the reform became effective on January 1st, 2004.

Table 2 shows the fiscal implications of the reform on DI spending by comparing the sum of DI benefits for the stock of DI beneficiaries in the year 2003 and applying once the old and once the new payout structure for individuals who were directly affected by the reform (pre-reform disability degree between 60 and 69% in 2003). Panel (A) of Table 2 presents the direct fiscal effect on spending assuming that individuals would not change their labor market behavior and keep the same disability level as in the pre-reform period. Without considering any behavioral change, the reform increased fiscal spending by about 73 million Swiss Francs. Panel (B) takes the behavioral change into account and compares spending from the new and the old payout structures but uses actually observed disability degrees for each year. The results show that realized DI savings for individuals who had initially a disability degree between 67 to 69% were about 13 million Swiss Francs lower than expected. Many of the concerned individuals thus managed to increase their disability degree

[1] Further features of the reform were the abolition of additional spousal pensions for new pensions and hardship pensions, the increase in the helpless allowance for individuals with special care needs and the development of medical expertise for the evaluation of DI benefit appraisals via the introduction of regional screening centers. These parts of the reform affect either only inflow or all insured individuals equally and are not tied to any age thresholds.

Table 2: Implication on total spendings for main DI pensions (2004-2007)

	2004 (1)	2005 (2)	2006 (3)	2007 (4)	2004-2007 (5)
PANEL A: EX-ANTE ANALYSIS ON THE STOCK OF 2003					
67-69%	-11.01	-11.06	-10.87	-10.99	-43.94
60-66%	31.44	30.56	28.46	27.27	117.74
Total	20.43	19.51	17.59	16.28	73.80
PANEL B: EX-POST ANALYSIS ON THE STOCK OF 2003					
67-69%	-8.40	-7.77	-7.23	-6.95	-30.35
60-66%	30.75	30.27	28.75	27.93	117.70
Total	22.35	22.51	21.51	20.98	87.35

Note: Numbers are in million CHF. The sample consists of all DI beneficiaries observed in the year 2003. Financial implications are predicted on spending for main benefits and do not consider additional benefits for children and spouses, helpless allowance or means tested benefits. Disability degrees in panel A relate to the ones observed in 2003. Panel B predicts the financial impact using observed disability degrees for the years 2004 to 2007.

and keep a full disability pension. In the following we are interested if this increase in the disability degree can be explained by work disincentives of the new payout structure.

3 Predicted effects in a revealed preferences framework

The expected effects of the reform can be predicted in a simple static labor supply model: Total disposable income y consists of earnings Y and disability benefits B. The DI suffers from an asymmetric information problem because DI caseworkers cannot observe the true disability degree. Assume that the caseworker sets potential earnings without disability equal to the last earnings before the onset of the disability, and potential earnings with disability equal to observed current earnings. The resulting thresholds in current earnings imposed by the step-wise DI system are thus unique for each individual and depend on their earnings before the onset of the disability. Individuals can signal a higher disability degree by choosing a lower employment level. Since our analysis focuses on individuals with disability degrees without the reform to be between 67 and 69%, we assume for simplicity one single notch:

$$y = \begin{cases} Y + B & \text{if} \quad Y \leq \pi \\ Y + \alpha B & \text{if} \quad Y > \pi \end{cases}$$

where $(1 - \alpha)B$ is the size of the notch and π is the earnings threshold for a full pension.

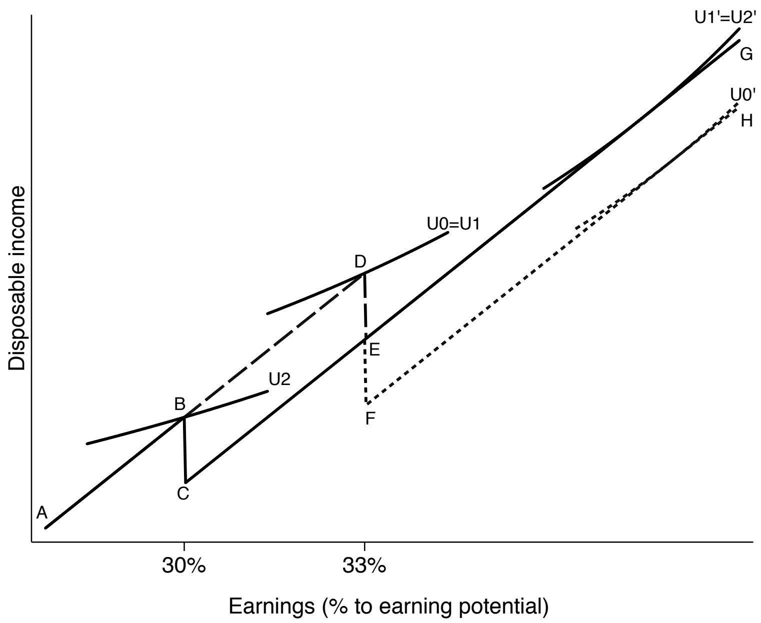
Figure 1 lays out the expected effect of the reform. Note that we label earnings on the abscissa in terms of a percentage to earnings potential to ease comparability with other settings, where fixed absolute earnings thresholds (such as the SGA threshold) are modified. Before the reform, the budget constraint is identical for all individuals in relative terms (ADFH). Individuals have the choice to sacrifice 50% of their disability pension and increase their earnings, or to reduce labor supply to 33% and receive the full DI pension. The individual chooses a full pension if the utility U of an employment level of 33% or less ($U0$) is higher than the utility from higher employment levels ($U0'$).

Individuals who are 50 years and older experience a parallel shift in the budget constraint for all earnings above 33%. The cut-off threshold for a full pension, however, stays constant at 33% since they are exempted from the benefit cut. Their budget constraint is thus equal to ADEG. They are now only sacrificing 25% of their DI benefit when employment exceeds 33%, which might cause some individuals to expand employment (if $U1 < U1'$). Note however, that only individuals who are cash cliff constrained (or in other words, bunch labor supply at the earnings threshold) will react to this aspect of the reform. Individuals who optimally chose labor supply to be less than 33% have no incentive to change behavior, since they are not affected by the discontinuity in the budget constraint.

Individuals who are younger than 50 years additionally experience a shift of the earnings threshold for a full pension from 33% to 30%. Their new budget constraint is ABCG. These individuals have two options: They either accept the pension cut, which results in an increase of labor supply due to the standard income effect (if $U2 < U2'$) or they reduce employment enough to fall below the new earnings threshold of 30% and keep their full pension due to the substitution effect (if $U2 \geq U2'$).

Moreover, shifting the cash cliff has no effect on older individuals who react to the reduction of the cash cliff to 25%. Note that for these individuals the preference structure is $U1 < U1'$ (as in the previous case). Since the slope of the budget constraint is not affected by the shift in the cash cliff, labor supply and utility remain unchanged $U1' = U2'$. Reducing employment to fall below the new earnings

Figure 1: Expected effect of the reform



Note: The line ADFH is the budget constraint for all individuals before the DI reform. The line ABCG is the budget constraint for individuals eligible for DI benefits and younger than 50 in January 2004. The line ADEG is the budget constraint for individuals older than 50 in January 2004. Disposable income is earnings plus DI pensions. Earnings are in percentage to earnings potential. Ui denotes utility levels if the person received a full pension, Ui' if the person receives a partial pension. The subscript i denotes the utility levels under different reforms (i.e. 0 without the reform, 1 for individuals older than 50 years, and 2 for individuals younger than 50).

threshold of 30% clearly yields lower utility than the status quo $U2 < U1$. Their revealed preference structure is therefore $U2 < U1 < U1' = U2'$.

The empirical part of this paper compares labor outcomes of individuals who were fully exposed to the reform to individuals who were exempted from the benefit cut. This means that we compare the difference between the solid and the dashed line in figure 1.

4 Identification

4.1 Total effect of the reform

In the first step we predict the total effect of the reform on employment and earnings. We make use of the fact that individuals aged 50 or older in January 2004 are exempted from the reform and thus serve as a control group.

Denote exposure to the reform with D_i, where $D_i = 1$ denotes the age cohort exposed to the reform and $D_i = 0$ denotes the age cohort not exposed. $Y_{it}(D_i)$ denotes potential outcome, where for each individual we can observe only $Y_{it} = D_i Y_{it}(1) + (1 - D_i)Y_{it}(0)$. The individual effect of the reform $TE_{it} = Y_{it}(1) - Y_{it}(0)$ cannot be identified, as we never observe the same individual in both potential outcome states. The simple comparison between average outcomes of treated and controls is biased by non-random treatment allocation (selection bias) because of age trends in health, employment, and earnings.[2]

We employ a difference-in-differences strategy to correct for this bias and estimate the average treatment effect on the treated (ATET) as follows:[3]

$$ATET_t = E[Y_t(1) - Y_t(0)|D = 1] = E[Y_t - Y_0|D = 1] - E[Y_t - Y_0|D = 0]$$

where the subscript t denotes the years after the introduction of the reform, and the subscript 0 refers to the last period before the reform (i.e. 2003). Since we have panel data, we implement this method by using a first-difference regression

$$\Delta Y_{it} = \alpha + \beta D_i + u_i, \tag{1}$$

where $\Delta Y_{it} = Y_{it} - Y_{i0}$, Y_{it} is the observed outcome for the years following the reform, Y_{i0} is the observed outcome for the year 2003, and β is the average treatment effect on the treated.[4]

The identification strategy relies on the parallel-trend assumption, or in other words the assumption that outcomes would develop parallel for the control and the treatment group in absence of the reform. To strengthen the validity of the parallel trend assumption we will focus on individuals in a relatively small bandwidth around the age threshold of fifty. In the limit, we employ a regression discontinuity design in first differences (FD-RDD) that estimates the local average treatment effect (LATE) directly at the age threshold of fifty. The identifying assumption is then, that treatment variation is locally randomized for individuals close to the age threshold. This assumption implies that mean potential outcomes without treatment

[2] Previous literature has exploited bunching to estimate behavioral responses to kinks or discontinuities in tax and benefit schedules (i.e. Brown, 2013; Ruh and Staubli, 2015; Saez, 2010). However, disability degrees are not smoothly distributed (see Figure A.1) and we observe strong bunching at decimal disability degrees that are not associated with payout thresholds (such as 70%, 80%, or 100%, for example).

[3] From now on, the subscript i will be omitted when possible.

[4] Note that this estimation procedure cannot take into account endogenous outflow. From 2003 to 2004, outflow in treated and control groups were low and comparable (2.7% vs. 2.2%).

are identical for treated and controls

$$\lim_{\epsilon \to 0} E(Y_t(0)|A = 50 - \epsilon) = \lim_{\epsilon \to 0} E(Y_t(0)|A = 50 + \epsilon),$$

where A denotes age on the first of January 2004. The parallel trend assumption thus holds by construction. The FD-RDD estimator can be implemented by estimating the following regression

$$\Delta Y_{it} = \alpha_0 + \beta_1 D_i + \beta_2 (A_i - 50) + \beta_3 (A_i - 50) D_i + u_i,$$

where β_1 measures the local average treatment effect.

4.2 Income and substitution effects

The total effect of the reform on employment and earnings is an average of conflicting income and substitution effects. We will use the principal stratification framework (Frangakis and Rubin, 2002) to decompose the total effect into group specific effects for different strata. Within our simple theoretical model the causal effects for the different principal strata have a straightforward economic interpretation as income and substitution effects. The following section outlines our empirical strategy. The exact formulation of the resulting bounds is summarized in table 3, while all mathematical proofs are provided in the appendix.

Denote potential partial DI pension receipt with an indicator $P_t(D) \in \{0, 1\}$, which is equal to one if a person receives a partial DI pension and equal to zero if the person receives a full pension. The indicator for potential partial DI pension receipt and exposure to the reform are both binary, allowing decomposing the full population into four different strata ($S_t = s$). For two groups, the pension level does not change due to the reform: Never-takers ($S_t = nt$) do never reduce pensions to partial benefits ($P_t(1) = 0, P_t(0) = 0$), while always-takers ($S_t = at$) always receive a partial pension ($P_t(1) = 1, P_t(0) = 1$). The remaining two strata react to the reform. Compliers ($S_t = c$) reduce pension level as a result of exposure to the reform ($P_t(1) = 1, P_t(0) = 0$), while defiers ($S_t = d$) show the exactly inverse reaction ($P_t(1) = 0, P_t(0) = 1$).

The principal effect with respect to a principal stratum is defined as the comparison of potential outcomes within a stratum. The average treatment effect on the treated can be decomposed to:

$$ATET_t = E[Y_t(1) - Y_t(0)|D = 1] \tag{2}$$

$$
\begin{aligned}
=&E[Y_t(1) - Y_t(0)|D = 1, S_t = nt]Pr(S_t = nt|D = 1) \\
&+ E[Y_t(1) - Y_t(0)|D = 1, S_t = c]Pr(S_t = c|D = 1) \\
&+ E[Y_t(1) - Y_t(0)|D = 1, S_t = d]Pr(S_t = d|D = 1) \\
&+ E[Y_t(1) - Y_t(0)|D = 1, S_t = at]Pr(S_t = at|D = 1)
\end{aligned}
$$

This decomposition brings the econometric model to the predictions of the simple labor supply model: To keep full pensions after the reform, never-takers need to increase their disability degree by signalling lower earnings potential. This may come at the cost of reducing labor supply so that earnings fall below the new earnings threshold. We therefore expect a negative earnings effect for never-takers due to the substitution effect $SE_t = E[Y_t(1) - Y_t(0)|D = 1, S_t = nt] \leq 0$. Compliers, in contrast, accept the pension-cut and potentially increase earnings due to the income effect $IE_t = E[Y_t(1) - Y_t(0)|D = 1, S_t = c] \geq 0$. Estimating principal effects thus yields the effects of interest for the two principal groups affected by the reform.

Analogue to the estimation of the total effect of the reform, we employ a difference-in-differences specification. Under the assumption of parallel trends within each strata **(Assumption 1.a)** we could estimate average treatment effects on the treated within each strata if the strata were observed. However, strata depend on latent variables and can therefore not be directly observed. Under the additional assumption that exposure to the reform is independent of potential pension receipt $P_{i,t}(1), P_{i,t}(0) \perp D_i$ **(Assumption 1.b)**, strata proportions are identical in the treated and control group.[5]

To partially identify substitution and income effects, we rely on an additional set of assumptions that is backed up by our simple labor supply model. Individual level monotonicity $P_{i,t}(1) \geq P_{i,t}(0)$ **(Assumption 2)** assures, that defiers do not exist. Nobody would decide to increase labor supply such as to receive only a partial pension in absence of the reform, but not increase labor supply to keep a full pension if affected by the reform. Individuals who choose a three-quarter pension over a full pension in absence of the reform reveal that their utility from expanding employment and lowering DI benefits is higher than the utility from bunching earnings at the old threshold, which is the necessary requirement to keep the full pension without the reform. Together, assumptions 1.b and 2 allow point identification of strata proportions p_s:

$$p_{nt} = Pr(P_t = 0|D = 1)$$

[5]Note that assumptions 1.a and 1.b impose parallel trends in the full sample, which was needed to identify the total effect of the reform.

$$
\begin{aligned}
p_{at} =& Pr(P_t = 1|D = 0) \\
p_c =& Pr(P_t = 1|D = 1) - Pr(P_t = 1|D = 0)
\end{aligned}
$$

These rather standard assumptions allow constructing bounds for the principal strata effects (equation 2). Since defiers are assumed not to exist (assumption 2), all treated individuals who keep a full pension are never-takers. We thus directly observe the first component of the first-difference estimator, i.e. $E[\Delta Y_t|D = 1, S_t = nt] = E[\Delta Y_t|D = 1, P_t = 0] = E[\Delta Y_t^{10}]$. However, never-takers and compliers are observationally equivalent in the observed group with $D = 0$ and $P_t = 0$. It is therefore not possible to directly observe the remaining components of the first difference estimator $E[\Delta Y_t|D = 0, S_t = nt]$. However, the relative group size of never-takers in the observed subgroup $D = 0$ and $P_t = 0$ can be estimated (i.e. $pr_{nt} = \frac{p_{nt}}{p_{nt}+p_c}$). To estimate the lower (upper) bound, we thus assign the largest (or smallest) values of ΔY_t for individuals with $D = 0$ and $P_t = 0$ to never-takers. Exactly the same approach can be used to bound the first-difference estimator for compliers. Here both components of equation 2 need to be bounded because compliers are observationally equivalent with never-takers in the observed group with $D = 0$ and $P_t = 0$, and with always-takers in the observed group with $D = 1$ and $P_t = 1$.

To further tighten these bounds, we can apply a set of additional assumptions that are predicted from our theoretical model:

Assumption 3: Exclusion restriction for always-takers

$$E[Y_t(0)|S_t = at] = E[Y_t(1)|S_t = at].$$

The exclusion restriction states that the reform has no effect on always-takers. This assumption is predicted by our theoretical model, since the relevant part of the budget constraint for individuals who choose a partial DI benefit in absence of the reform is the one to the right of the old notch. Neither the intercept nor the slope of this part of the budget constraint is affected by the reform. These individuals have thus no incentive to change labor supply at the intensive margin.[6] This assumption is closely related to the standard IV approach (Angrist et al., 1996; Imbens and Angrist, 1994). The key difference is, however, that we apply the exclusion restriction only to always-takers but not to never takers. Assumption 3 tightens upper and lower bounds for the complier population since they are observationally equivalent to

[6]The reform changed the intercept of the budget constraint, but it did so for individuals older and younger than fifty equally. Formally it must hold that $U1' = U2'$, see section 3.

Table 3: Bounds for income and substitution effect

Ass.	Lower bound	Upper bound
	Compliers: $S_t = c$	
1.a, 1.b, 2	$E[\Delta Y_t^{11} \vert \Delta Y_t^{11} \leq \Delta y_{t,pr_c}^{11}] - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} > \Delta y_{t,pr_{nt}}^{00}]$	$E[\Delta Y_t^{11} \vert \Delta Y_t^{11} > \Delta y_{t,pr_{at}}^{11}] - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} \leq \Delta y_{t,pr_c}^{00}]$
+ 3	$E[\Delta Y_t^{11}] - \frac{p_{at}}{p_c}\{E[\Delta Y_t^{11}] - E[\Delta Y_t^{01}]\} - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} > \Delta y_{t,pr_{nt}}^{00}]$	$E[\Delta Y_t^{11}] - \frac{p_{at}}{p_c}\{E[\Delta Y_t^{11}] - E[\Delta Y_t^{01}]\} - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} \leq \Delta y_{t,pr_c}^{00}]$
+ 4	$\frac{ATET_t}{p_c}$	$E[\Delta Y_t^{11}] - \frac{p_{at}}{p_c}\{E[\Delta Y_t^{11}] - E[\Delta Y_t^{01}]\} - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} \leq \Delta y_{t,pr_c}^{00}]$
	Never-takers: $S_t = nt$	
1.a, 1.b, 2	$E[\Delta Y_t^{10}] - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} > \Delta y_{t,pr_c}^{00}]$	$E[\Delta Y_t^{10}] - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} \leq \Delta y_{t,pr_{nt}}^{00}]$
+ 3	$E[\Delta Y_t^{10}] - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} > \Delta y_{t,pr_c}^{00}]$	$E[\Delta Y_t^{10}] - E[\Delta Y_t^{00} \vert \Delta_t^{00} \leq \Delta y_{t,pr_{nt}}^{00}]$
+ 4	$E[\Delta Y_t^{10}] - E[\Delta Y_t^{00} \vert \Delta Y_t^{00} > \Delta y_{t,pr_c}^{00}]$	0

Note: ΔY_t^{dp} denotes first difference of outcomes in the observed group with $D = d$ and $P_t = p$. $\Delta y_{t,pr_s}^{dp}$ denotes the pr_s-quantile of ΔY_t, where pr_s is equal to the relative size of stratum $S_t = s$ within the observed group with $D = d$ and $P_t = p$.

always-takers in the observed group with $D = 1$ and $P_t = 1$.

Moreover, assumptions 2 and 3 imply that the total treatment effect is equal to the weighted average of income and substitution effects:

$$\begin{aligned} ATET_t =& E[Y_t(1) - Y_t(0)|D = 1] \\ =& E[Y_t(1) - Y_t(0)|D = 1, S_t = nt]Prob(S_t = nt|D = 1) \\ &+ E[Y_t(1) - Y_t(0)|D = 1, S_t = c]Prob(S_t = c|D = 1) \end{aligned}$$

Assumption 4: Weak monotonicity of mean potential outcomes within strata

$$E[\Delta Y_t|D = 1, S_t = nt] \leq E[\Delta Y_t|D = 0, S_t = nt]$$
$$E[\Delta Y_t|D = 1, S_t = c] \geq E[\Delta Y_t|D = 0, S_t = c]$$

The model predicts a negative substitution effect and a positive income effect. This assumption implies that the upper bound for the principal effect for never-takers is equal to zero (since our estimate for the total effect is positive, see section 6.1), while the lower bound for the effect in the complier group becomes the standard Wald estimator.[7] Table 3 summarizes the bounds, depending on the imposed assumptions.

5 Data and descriptive statistics

The analysis is based on administrative data of the full sample of DI beneficiaries in Switzerland. We observe employment and earnings, DI pensions, the disability degree, and background characteristics (age, type of disability, canton of residence, marital status, citizenship, gender) in December of each year. The empirical analysis follows the stock of DI beneficiaries of the year 2003 up to 2007. Individuals having congenital disorders are excluded since special rules apply for determining their disability degree and pension size. The focus of this paper is on individuals who had a disability degree between 67 and 69% in 2003. This is only a small proportion of the full sample (3.4%).

Identification of the structural labor supply parameters relies on a cohort discontinuity, where individuals who were 50 years and older when the reform came into effect are exempted from the benefit cut and thus serve as control group. This is a suitable control group as it is not possible to manipulate age. However, we focus on

[7] In Appendix A.2 we also discuss the general case how this assumption affects bounds if the ATET is negative.

a selected sample and the reform was already discussed in 2002. Anticipation effects may thus lead to a situation, where individuals self-select in or out of the sample based on their age. For example, individuals who were younger than 50 years may have anticipated the reform and selected themselves out of the sample by adjusting their labor supply accordingly. Figure A.2 in appendix A.3 shows that this is not a major issue. There is no discontinuity in the age distribution among individuals with disability degrees between 67% and 69% in 2003.[8]

In the main empirical analysis we restrict the sample to individuals aged 42 to 57 in January 2004. Table 4 shows descriptive statistics for DI beneficiaries in the treatment and control group for the year 2003. As the treatment group contains younger individuals, it is not surprising that it is characterized by higher earnings and a higher probability to work. While the disability degree is by construction very similar for the two groups, the treatment group is characterized by slightly lower earnings before disability, and a higher probability to have a mental illness or an accident as a reason for DI, as opposed to musculoskeletal diseases. Other background characteristics are well balanced between the two groups.

The main outcome variables are employment (equal to one if the person has any earnings during the year) and yearly earnings. Trends in these outcomes and in the disability degree for the years 2001 to 2007 are presented in figure 2. Earnings (panel a) and labor supply (panel b) are higher for treated compared to controls, but these variables follow a parallel development before the reform came into effect in 2004. After the reform, the gap between treated and control seems to widen slightly. The disability degree (panel c) is by construction equal in 2003, but rises strongly after the reform for the treated group while it follows a smooth time trend for the control group.

6 Results

6.1 Total effect of the reform

Estimates for the total effect for the year 2004 are presented in table 5. Columns (1) and (2) show difference-in-differences estimates without and with the inclusion of background characteristics, respectively. Columns (3) and (4) display regression discontinuity estimates, where the effect of the reform is measured directly at the

[8] We explore anticipation using individuals with disability degrees between 67% and 69% in 2001 (or in other words before the reform was first discussed) in a robustness analysis.

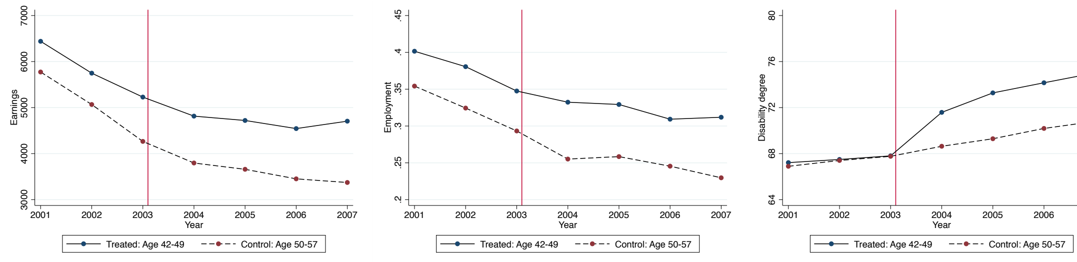
Table 4: Descriptive statistics for December 2003

	Treated (1)	Control (2)
PANEL A: OUTCOMES		
Annual earnings (CHF)	5228	4265
	(10363)	(9138)
Employment	0.35	0.29
	(0.48)	(0.46)
Disability degree	67.8	67.8
	(0.82)	(0.80)
Total pension (CHF)	27923	26143
	(12049)	(9741)
PANEL B: INDEPENDENT VARIABLES		
Avg. earnings during contribution time	47514	49622
	(26625)	(23868)
Mental illness	0.29	0.23
	(0.45)	(0.42)
Musculoskeletal disease	0.29	0.38
	(0.45)	(0.49)
Accident	0.20	0.15
	(0.40)	(0.36)
Age	45.9	53.9
	(2.3)	(2.3)
Married	0.64	0.68
	(0.48)	(0.47)
Foreigner	0.33	0.32
	(0.47)	(0.47)
Female	0.47	0.48
	(0.50)	(0.50)
Observations	1364	2305

Note: Standard deviations in parentheses. earnings are equal to zero if the individual is not employed. Employment is a dummy variable equal to 1 if earnings are larger than zero. Disability degree denotes earnings loss with disability as a percentage of potential earnings without disability. Total pension includes main pension, child pension, spouse pension, means tested benefits and helpless allowance. Information on type of disability is only available for a subset of all observations (1,335 treated and 2,252 controls).

Figure 2: Time trends for treated and controls

(a) Earnings (b)Employment (c)Disability degree



Note: Solid lines represent yearly mean outcomes for treated, dashed lines for control individuals. The sample consists of individuals aged between 42 and 57 in January 2004 and having a disability degree between 67 and 69% in December 2003. Outcomes are measured in December of each year. Vertical lines represent the time of the reform (January 2004).

age threshold of 50.

Panel A presents the total effect of the reform on earnings. The dependent variable is the change in earnings between 2004 and 2003. Column (1) shows that average earnings decreased over time in both groups. The control group (represented by the constant) decreased earnings by 453 CHF. Independent of the estimation method, the average effect of the reform on earnings is positive but very small and not significantly different from zero. Panel B focuses on employment. The causal effect of the reform amounts to 2.3 percentage points increase in employment, which is not negligible given the baseline employment share of 35 percent in the treated population. The FD-RDD estimator yields similar results, even though somewhat smaller and no longer statistically significant.

Panel C shows that the reform leads to a large increase in the average disability degree. The disability degree of the treated group increases by around 3 percentage points, which is just enough to reach the new threshold level of a full pension.

One might argue that particularly labor supply effects need longer time horizons to materialize. Table A.2 presents the effects for the years 2005 to 2007. Dependent variables are differences with respect to the base year 2003 ($\Delta Y_{it} = Y_{it} - Y_{i0}$). Estimates for all years are comparable to the immediate impact. We find only small and in most cases insignificant impacts on employment or earnings, but very stable and persistent effects on the disability degree and benefits. It thus seems that the reform of 2004 acted as a shock that had an immediate impact on how the disability insurance assessed the earnings potential of the insured, but had no short or long-term impact on average labor market outcomes.

The main estimates are subjected to several specification and robustness checks.

Table 5: Total effect of the reform

	DiD		FD-RDD	
	No controls (1)	Controls (2)	No controls (3)	Controls (4)
	PANEL A: EARNINGS			
Treated	33.5	29.0	134.0	161.6
	(173.2)	(174.6)	(313.5)	(319.3)
Constant	-452.7***	-24.5	-440.1*	-21.5
	(102.2)	(482.4)	(232.4)	(536.8)
R-squared	0.000	0.011	0.000	0.011
	PANEL B: EMPLOYMENT			
Treated	0.023**	0.023**	0.014	0.016
	(0.009)	(0.009)	(0.018)	(0.018)
Constant	-0.036***	-0.031	-0.032**	-0.028
	(0.005)	(0.025)	(0.013)	(0.027)
R-squared	0.002	0.009	0.002	0.009
	PANEL C: DISABILITY DEGREE			
Treated	2.910***	2.887***	3.180***	3.162***
	(0.275)	(0.271)	(0.493)	(0.488)
Constant	0.874***	1.821	0.727***	1.762
	(0.103)	(1.403)	(0.238)	(1.458)
R-squared	0.040	0.069	0.041	0.070
Observations	3,581	3,581	3,581	3,581

Note: Robust standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Sample consist of individuals aged between 42 and 57 in January 2004 and having a disability degree between 67 and 69% in December 2003. Outcomes are first differences between the years 2004 and 2003. Earnings are yearly earnings in CHF. Earnings are set to zero if not working. Employment is a dummy variable equal to 1 if the individual has earnings above zero. Disability degree is potential earnings loss as a percentage of potential earnings without disability. Controls include dummies for canton of residence and year of observation, marital status, dummy for Swiss citizenship and gender.

Table A.3 in the appendix A.3 shows that the estimates are robust to different choices of age bandwidths and the inclusion of squared terms. Table A.4 uses the year 2001 as baseline and focusses on the stock of 2001 with disability degrees between 67 and 69% in 2001 as the reference population. Results are very similar to the main specification suggesting that anticipation does not bias the results.[9]

6.2 Income and substitution effects

We now apply the bounds derived in section 4 to predict income and substitution effects. Table 6 shows the respective strata proportions in column (1). This is equivalent to the cross-sectional estimator, where a binary indicator for partial pension receipt is used as dependent variable. The share of always-takers (represented by the constant) is with 0.6% very low. This means that in absence of the reform, hardly anybody increases labor supply above the earnings threshold and receives a partial DI pension.[10] The reform increases the share of individuals who receive a partial pension by 24.6 percentage points. These individuals are thus compliers to the reform. The remaining share (74.8%) are then never-takers. Controlling for background characteristics and/or estimating the strata proportions directly at the age threshold using a RDD regression yields comparable results.

Table 7 shows the estimated bounds for the income and substitution effects. Bounds imposing only parallel trends, independence of strata across treatment groups and monotonicity are wide and include the zero. The estimated treatment effect on yearly earnings for compliers reaches from -2,864 to 3,730 CHF. Bounds for never-takers are tighter, reaching from -1,001 to 861 CHF. Assuming the exclusion restriction for always-takers allows to tighten the bounds for compliers only slightly to -2,513 to 3,135 CHF, leaving bounds for never-takers unaffected. Assuming weak monotonicity of mean potential outcomes within strata (assumption 4) implies that the upper bound for the principal effect for never-takers must be equal to zero since the total effect is slighly positive, and that the lower bound for the compliers becomes the standard Wald estimator. This leads to informative bounds for the compliers

[9] Table A.5 shows placebo tests. When the first-difference estimator is applied, we observe some statistically significant coefficients, but not more/less than what would be expected by chance. FD-RDD estimates are small and insignificant in all cases.

[10] These results implicitly confirm the findings from Bütler et al. (2015), who study the effect of a conditional cash program that is paid out to DI beneficiaries if they increase employment and lower DI benefits by at least one quarter. Recall from section 3, that always-takers are individuals who expanded employment because the introduction of the three-quarter pension reduces the cash-cliff for the next lower pension level by 25 percentage points ($U0 = U1 < U1' = U2'$). The very low share of always-taker in our data is quantitatively very similar to the low take-up rate of the conditional cash payment of 0.5% documented by Bütler et al. (2015).

Table 6: Estimation of strata proportions

	Diff. in means		RDD	
	No controls (1)	Controls (2)	No controls (3)	Controls (4)
Treated	0.246***	0.249***	0.193***	0.197***
	(0.012)	(0.012)	(0.020)	(0.019)
Constant	0.006***	-0.059	0.007*	-0.069*
	(0.002)	(0.039)	(0.004)	(0.039)
R-squared	0.161	0.184	0.167	0.190
Observations	3,581	3,581	3,581	3,581

Note: Robust standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Sample consist of individuals aged between 42 and 57 in January 2004 and having a disability degree between 67 and 69% in December 2003. Outcome is a dummy for partial pension receipt in 2004. Controls include dummies for canton of residence and year of observation, marital status, dummy for Swiss citizenship and gender.

that range from 136 to 3,135 CHF.

Bounds for employment effects are informative for compliers but also very large. The reform has a positive impact on labor market participation for compliers of between 9.3 and 20.8 percentage points. This is a substantial increase compared to mean pre-reform labor market participation of 35%. For never-takers bounds cover the zero and are much smaller. If anything, never-takers reduce labor supply on the extensive margin by at most 3.7 percentage points.

We also predict bounds for principal effects on the disability degree. Since the disability degree denotes the expected earnings loss due to a disability in relative terms, we expect a reverse sign for principal effects.[11] We find tight bounds for never-takers suggesting that disability degrees increase by 4 to 5 percentage points. For compliers, bounds for the principal effects are somewhat larger ranging from -1 to -5 percentage points.

The results provide evidence that particularly the income effects play a significant role in explaining high dependence to social assistance and low labor supply among the disabled. On the other hand, the relatively small substitution effects also demonstrate the limitation when applying the standard labor market models to the disability insurance: People managed to keep a full pension without the need to

[11] When deriving bounds for the effect on the disability degree, we need to take into account that always-takers decrease their disability degrees below 66% to receive a partial pension, which affects bounds that are predicted employing assumption 3 (exclusion restriction for always takers).

Table 7: Bounds for income and substitution effect

	Earnings		Employment		Disability degree	
Ass.	Lower bound (1)	Upper bound (2)	Lower bound (3)	Upper bound (4)	Lower bound (5)	Upper bound (6)
			COMPLIERS: $S_t = c$			
1.a, 1.b, 2	-2864***	3730***	-0.090***	0.230***	-11.58***	-0.89***
	(427)	(513)	(0.023)	(0.031)	(0.53)	(0.22)
+ 3	-2513***	3135***	-0.060***	0.208***	-4.86***	-1.07***
	(428)	(506)	(0.022)	(0.030)	(0.53)	(0.22)
+ 4	136	3135***	0.093***	0.208***	-4.86***	-1.07***
	(500)	(506)	(0.030)	(0.030)	(0.53)	(0.22)
			NEVER-TAKERS: $S_t = nt$			
1.a, 1.b, 2	-1001***	861***	-0.037***	0.052***	4.17***	5.42***
	(207)	(216)	(0.010)	(0.012)	(0.39)	(0.36)
+ 3	-1001***	861***	-0.037***	0.052***	4.17***	5.42***
	(207)	(216)	(0.010)	(0.012)	(0.39)	(0.36)
+ 4	-1001***	0	-0.037***	0	4.17***	5.42***
	(207)		(0.010)		(0.39)	(0.36)

Note: Bootstrapped standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Sample consist of individuals aged between 42 and 57 in January 2004 and having a disability degree between 67 and 69% in December 2003. Outcomes are first differences between the years 2004 and 2003. Employment is a dummy variable equal to 1 if the individual has earnings above zero. Earnings are yearly earnings in CHF. Earnings are set to zero if not working. Disability degree is potential earnings loss as a percentage of potential earnings without disability. Calculation of bounds for earnings and employment as shown in table 3. Calculation of bounds for disability degree take into account that always-takers need to have a disability degree below 67%.

signal over a labor supply response.

This can be demonstrated by a simple back-of-the envelope calculation: DI caseworkers typically set potential earnings without a disability equal to the last earnings before the onset of the disability. This variable is unknown but can be proxied by the average earnings during contributing years before the onset of the disability. Table A.6 shows that never-takers had average annual earnings before the onset of disability of 46,939 Swiss Francs. If the increase in the average disability degree by 4.17 to 5.42 percentage points would be fully driven by labor market responses, the substitution effect should amount to between -1,957 and -2,544 Swiss Francs. Our estimated substitution effect, however, is far lower, which indicates that at least half of the increase in the average disability degree is indeed not driven by the substitution effect. These results provide evidence that actual labor market behavior is in many cases not used to assess the disability degree, and they show also the limitations of the process determining the disability degree, which is consistent with other empirical evidence based on the Swiss disability insurance (Liebert, 2016).

7 Conclusion

This paper evaluates a reform of the Swiss disability insurance system that introduced the three-quarter pension and thus further graduated the existing partial system. The main analysis focuses on those DI beneficiaries who, with their pre-reform disability degree, would lose a quarter pension and are faced with a lower earnings threshold to remain eligible for the full pension. We find that the reform reduced average DI benefits from the public DI system on average by CHF 1,800 per year (ca. 1,450 USD or 1,150 EUR based on exchange rates when the reform came into force), which represents a 7% reduction of average DI benefits. Effects on average labor supply on the intensive and extensive margin are modest, but consist of conflicting income and substitution effects. We partially identify these effects. Bounds for income effects suggest that individuals who complied to the reform (ca. 25%) increased yearly earnings between 136 to 3,135 CHF. This is a sizeable amount compared to previous earnings (max 50% increase in earnings). Bounds for the substitution effect imply that individuals who kept a full pension reduced yearly earnings by a maximum of CHF 1,000 (if any).

There is a huge public interest in reforms that remove existing work disincentives caused by cash cliffs. A widely discussed reform is to further graduate the DI benefit payout structure. Policy makers need to be aware that this policy can lead to conflicting income and substitution effects, making the total effect ambiguous.

In a nutshell, such policies should be only implemented if the substitution effect is severe and many DI beneficiaries are cash cliff constrained. Our paper shows, that the substitution effect is not the driving force imposing work disincentives to DI beneficiaries.

However, the results call attention to two aspects: First, the replacement rates guaranteed by different non-earned income sources are very high in Switzerland. The income effect may therefore be the driving channel leading to low labor supply and high dependency on DI benefits among the disabled. Second, linking the pay-out structure directly to disability induced income losses is a doubtful concept. The strong impact of the reform on disability degrees suggests that the disability degree can be manipulated. We show that DI beneficiaries signalling a higher earnings loss by choosing lower labor supply can be, if at all, only partially blamed. The available empirical evidence rather implies that the disability insurance reassessed the disability degree in favor of the insured individuals without the need of a corresponding labor supply response. This means that many DI beneficiaries were not forced to reduce labor supply, but were rather reclassified by the disability insurance.

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A Appendix

A.1 Point identification of strata proportions

Principal strata, and the mix of principal strata observed in groups with values $P_t = p$ and $D = d$ are presented in table A.1. **Assumption 2 (monotonicity)** rules out that defiers exist. We can therefore directly observe the population proportion of never-takers and always-takers in treated and control groups:

$$p_{nt|D=1} = Pr(P_t = 0|D = 1)$$

$$p_{at|D=0} = Pr(P_t = 1|D = 0)$$

Assumption 1.b (independence of potential pension receipt) states that the potential pension level, and thus principal strata, is independent of exposure to the reform:

$$p_{nt} = p_{nt|D=1} = p_{nt|D=0}$$

$$p_{at} = p_{at|D=1} = p_{at|D=0}$$

This allows to point identify the population proportion of compliers:

$$p_c = Pr(P_t = 1|D = 1) - Pr(P_t = 1|D = 0)$$

Table A.1: Types of individuals in potential treatment and observed treatment

Potential pension level			Observed groups		
$P_t(1)$	$P_t(0)$		P_t	D	
0	0	Never-taker	1	1	Complier or Always-taker
1	0	Complier	0	0	Complier or Never-taker
0	1	Defier	0	1	Defier or Never-taker
1	1	Always-taker	1	0	Defier or Always-taker

A.2 Partial identification of principal effects

In this paper we seek to estimate the causal effect of exposure to the reform on labor supply in different latent groups. **Assumption 1.a (parallel trends within strata)** allows to rewrite the effect of interest as the first-difference estimator in stratified samples:

$$E[Y(1) - Y(0)|D = 1, S_t = s] = E[\Delta Y_t|D = 1, S_t = s] - E[\Delta Y_t|D = 0, S_t = s]$$

Under **assumption 2 (monotonicity)**, we can directly observe the first component of the first-difference estimator for compliers ($E[\Delta Y_t|D = 1, S_t = nt] = E[\Delta Y_t|D = 1, P_t = 0] = E[\Delta Y_t^{10}]$) and the second component for always-takers ($E[\Delta Y_t|D = 0, S_t = at] = E[\Delta Y_t|D = 0, P_t = 1] = E[\Delta Y_t^{01}]$). The remaining components of the first-difference estimators cannot be observed, because compliers are observationally equivalent with always-takers in the group with $P_t = 1$ and $D = 1$, and with never-takers in the group with $P_t = 0$ and $D = 0$, respectively.

We construct worst-case scenarios for the remaining unobserved components by recognizing that average ΔY_t for individuals in mixed observed groups (such as $P_t = 1$ and $D = 1$ for example) can be written as

$$E[\Delta Y_t^{11}] = \frac{p_{at}}{p_{at} + p_c} E[\Delta Y_t|D = 1, S_t = at] + \frac{p_c}{p_{at} + p_c} E[\Delta Y_t|D = 1, S_t = c]$$

Since strata proportions are point identified, $E[\Delta Y_t|D = 1, S_t = at]$ can be bounded from above (below) by the the $pr_{at} = \frac{p_{at}}{p_{at}+p_c}$ fraction of the largest (smallest) values of ΔY_t for individuals in the observed group with $P_t = 1$ and $D = 1$. The resulting worst-case bounds for always-taker is than equal to

$$LB_{at}^{wc} = E[\Delta Y_t^{11}|\Delta Y_t^{11} \leq \Delta y_{t,pr_{at}}^{11}] - E[\Delta Y_t^{01}]$$

$$UB_{at}^{wc} = E[\Delta Y_t^{11}|\Delta Y_t^{11} > \Delta y_{t,pr_c}^{11}] - E[\Delta Y_t^{01}],$$

where $\Delta y_{t,pr_s}^{pd}$ denotes the pr_s quantile of ΔY_t in the group $P_t = p$ and $D = d$.

We follow the same approach to bound the principal effect for never-taker:

$$LB_{nt}^{wc} = E[\Delta Y_t^{10}] - E[\Delta Y_t^{00}|\Delta Y_t^{00} > \Delta y_{t,pr_c}^{00}]$$

$$UB_{nt}^{wc} = E[\Delta Y_t^{10}] - E[\Delta Y_t^{00}|\Delta Y_t^{00} \leq \Delta y_{t,pr_{nt}}^{00}]$$

For compliers we take into account, that both components of the first difference estimator needs to be bounded in the same manner:

$$LB_c^{wc} = E[\Delta Y_t^{11}|\Delta Y_t^{11} \leq \Delta y_{t,pr_c}^{11}] - E[\Delta Y_t^{00}|\Delta Y_t^{00} > \Delta y_{t,pr_{nt}}^{00}]$$

$$UB_c^{wc} = E[\Delta Y_t^{11}|\Delta Y_t^{11} > \Delta y_{t,pr_{at}}^{11}] - E[\Delta Y_t^{00}|\Delta Y_t^{00} \leq \Delta y_{t,pr_c}^{00}]$$

Assumption 3 (exclusion restriction) states that the principal effect for always-takers is zero, which point identifies the first component of the treatment effect for always-takers

$$E[\Delta Y_t|D=1, S_t=at] = E[\Delta Y_t|D=0, S_t=at] = E[\Delta Y_t^{01}]$$

This allows to point identify the first component of the first-difference estimator for compliers via the observation that the observed ΔY_t in the group with $P_t=1$ and $D=1$ is the weighted average of ΔY_t for compliers and always-takers:

$$E[\Delta Y_t|D=1, S_t=c] = E[\Delta Y_t^{11}] - \frac{p_{at}}{p_c}\{E[\Delta Y_t^{11}] - E[\Delta Y_t^{01}]\}$$

Bounds for compliers are thus tightened to

$$LB_c^{+A3} = E[\Delta Y_t^{11}] - \frac{p_{at}}{p_c}\{E[\Delta Y_t^{11}] - E[\Delta Y_t^{01}]\} - E[\Delta Y_t^{00}|\Delta Y_t^{00} > \Delta y_{t,pr_{nt}}^{00}]$$

$$UB_c^{+A3} = E[\Delta Y_t^{11}] - \frac{p_{at}}{p_c}\{E[\Delta Y_t^{11}] - E[\Delta Y_t^{01}]\} - E[\Delta Y_t^{00}|\Delta Y_t^{00} \leq \Delta y_{t,pr_c}^{00}]$$

Assumption 3 does not affect the bounds for never-takers, because they never share an observed group with always-takers

$$LB_{nt}^{+A3} = LB_{nt}^{+wc}$$

$$UB_{nt}^{+A3} = UB_{nt}^{+wc}$$

Assumption 3 furthermore allows rewriting the total effect into a weighted average of principal stratum effects for compliers and never-takers:

$$\begin{aligned}
ATET_t =& E[Y_t(1) - Y_t(0)|D=1] \\
=& E[Y_t(1) - Y_t(0)|D=1, S_t=nt]Prob(S_t=nt) \\
&+ E[Y_t(1) - Y_t(0)|D=1, S_t=c]Prob(S_t=c) \\
=& \{E[\Delta Y_t|D=1, S_t=nt] - E[\Delta Y_t|D=0, S_t=nt]\}p_{nt} \\
&+ \{E[\Delta Y_t|D=1, S_t=c] - E[\Delta Y_t|D=0, S_t=c]\}p_c
\end{aligned}$$

Assumption 4 (weak monotonicity) predicts the sign of the principal stratum effects, i.e. $E[Y(1){-}Y(0)|D=1, S_t=nt] \leq 0$ and $E[Y(1){-}Y(0)|D=1, S_t=c] \geq 0$. This means that either the upper bound for never-takers and/or the lower bound for compliers is equal to zero, depending on the sign of the $ATET_t$. Via the observation that the total effect is equal to the weighted average of principal stratum effects, the corresponding upper (lower) bound for complier (never-taker) is equivalent to the standard Wald estimator, i.e. the average effect of the reform in the full population weighted by the inverse of the probability that a person belongs to the respective stratum. In case that $ATET_t = 0$, lower bounds for compliers and upper bound for never-takers are both zero. Upper bounds for compliers and lower bounds for never-takers, however, are not affected. For compliers, bounds thus become:

$$LB_c^{+A4} = \max\left(0, \frac{\mathrm{ATET_t}}{p_c}\right)$$

$$UB_c^{+A4} = UB_c^{+A3}$$

Bounds for never-takers are:

$$LB_{nt}^{+A4} = UB_{nt}^{+A3}$$

$$UB_{nt}^{+A4} = \min\left(0, \frac{\mathrm{ATET_t}}{p_{nt}}\right)$$

A.3 Supplementary tables and figures

Table A.2: Long run effects

	DiD			RDD		
	2005	2006	2007	2005	2006	2007
	(1)	(2)	(3)	(4)	(5)	(6)
			Panel A: Earnings			
Treated	62.8	70.5	357.0	523.2	265.3	481.0
	(212.7)	(225.6)	(239.9)	(413.9)	(437.3)	(496.7)
R-squared	0.011	0.012	0.011	0.011	0.012	0.012
			Panel B: Employment			
Treated	0.018	0.008	0.023*	0.025	-0.005	0.006
	(0.011)	(0.012)	(0.013)	(0.021)	(0.023)	(0.024)
R-squared	0.011	0.006	0.009	0.011	0.007	0.009
			Panel C: Disability degree			
Treated	3.878***	3.845***	4.034***	4.104***	3.936***	4.088***
	(0.328)	(0.359)	(0.384)	(0.607)	(0.683)	(0.739)
R-squared	0.087	0.082	0.088	0.087	0.082	0.088
Observations	3,497	3,426	3,363	3,497	3,426	3,363

Robust standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Sample consist of individuals aged between 42 and 57 in January 2004 and having a disability degree between 67 and 69% in December 2003. Outcomes are first differences between the years 2005, 2006, 2007 and 2003, respectively. Earnings are yearly earnings in CHF. Earnings are set to zero if not working. Employment is a dummy variable equal to 1 if the individual has earnings above zero. Disability degree is potential earnings loss as a percentage of potential earnings without disability. Total pension includes main pension, child pension, spousal pension, means tested benefits, and helpless allowances. Controls include dummies for canton of residence and year of observation, marital status, dummy for Swiss citizenship and gender.

Table A.3: Specification checks: Age bandwidth and functional form

	DiD		RDD		
	44-55	40-59	44-55	40-59	squares
	(1)	(2)	(3)	(4)	(5)
			PANEL A: EARNINGS		
Treated	92.7	98.4	180.4	-95.5	165.4
	(200.9)	(167.3)	(352.2)	(312.2)	(473.3)
R-squared	0.019	0.009	0.019	0.010	0.012
			PANEL B: EMPLOYMENT		
Treated	0.027**	0.025***	0.000	0.015	-0.024
	(0.011)	(0.009)	(0.021)	(0.016)	(0.029)
R-squared	0.012	0.007	0.013	0.007	0.010
			PANEL C: DISABILITY DEGREE		
Treated	3.076***	2.945***	2.771***	2.996***	2.829***
	(0.305)	(0.250)	(0.534)	(0.444)	(0.693)
R-squared	0.080	0.068	0.080	0.068	0.070
Observations	2,623	4,496	2,623	4,496	3,581

Note: Robust standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Sample consist of individuals aged between 44 and 55 in January 2004 (columns 1 and 3), aged between 40 and 59 (columns 2 and 4), aged between 42 and 57 (column 5) and having disability degree between 67 and 69% in December 2003. Outcomes are first differences between the years 2004 and 2003. Earnings are yearly earnings in CHF. Earnings are set to zero if not working. Employment is a dummy variable equal to 1 if the individual has earnings above zero. Disability degree is potential earnings loss as a percentage of potential earnings without disability. Total pension includes main pension, child pension, spousal pension, means tested benefits, and helpless allowances. Controls include dummies for canton of residence and year of observation, marital status, dummy for Swiss citizenship and gender.

Table A.4: Sensitivity check: Base year 2001

	DiD	RDD
	(1)	(2)
PANEL A: EARNINGS		
Treated	450.8	-358.6
	(306.3)	(606.4)
R-squared	0.016	0.017
PANEL A: EMPLOYMENT		
Treated	0.017	-0.014
	(0.015)	(0.028)
R-squared	0.008	0.009
PANEL C: DISABILITY DEGREE		
Treated	2.633***	2.916***
	(0.383)	(0.721)
R-squared	0.055	0.055
Observations	2,669	2,669

Robust standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Sample consist of individuals aged between 42 and 57 in January 2004 and having a disability degree between 67 and 69% in December 2001. Outcomes are first differences between the years 2004 and 2001. Earnings are yearly earnings in CHF. Earnings are set to zero if not working. Employment is a dummy variable equal to 1 if the individual has earnings above zero. Disability degree is potential earnings loss as a percentage of potential earnings without disability. Total pension includes main pension, child pension, spousal pension, means tested benefits, and helpless allowances. Controls include dummies for canton of residence and year of observation, marital status, dummy for Swiss citizenship and gender.

Table A.5: Placebo tests: Anticipation, age, disability degree

	Anticipation	Age		Disability degree	
	2003	34-49	50-65	60-65	70-75
	(1)	(2)	(3)	(4)	(5)
			Panel A: Earnings		
DiD	77.2	88.6	206.1	979.1***	132.2
	(266.6)	(235.2)	(191.6)	(308.0)	(97.8)
R-squared	0.014	0.034	0.006	0.015	0.005
FD-RDD	-515.8	124.9	545.9	230.0	-144.5
	(564.2)	(478.8)	(446.7)	(654.6)	(221.1)
R-squared	0.015	0.034	0.007	0.015	0.006
			Panel B: Employment		
DiD	-0.009	-0.009	-0.001	0.008	0.005
	(0.013)	(0.013)	(0.008)	(0.010)	(0.005)
R-squared	0.011	0.007	0.008	0.014	0.001
FD-RDD	-0.023	0.001	0.014	-0.019	-0.011
	(0.025)	(0.026)	(0.016)	(0.021)	(0.009)
R-squared	0.012	0.008	0.008	0.014	0.002
			Panel C: Disability degree		
DiD	0.163	0.096	0.223	-0.605**	0.062
	(0.204)	(0.433)	(0.140)	(0.276)	(0.053)
R-squared	0.032	0.055	0.022	0.013	0.004
FD-RDD	-0.193	1.262	0.212	0.329	0.029
	(0.418)	(0.857)	(0.266)	(0.548)	(0.105)
R-squared	0.032	0.056	0.022	0.014	0.004
Observations	3,089	2,013	4,185	3,694	13,642

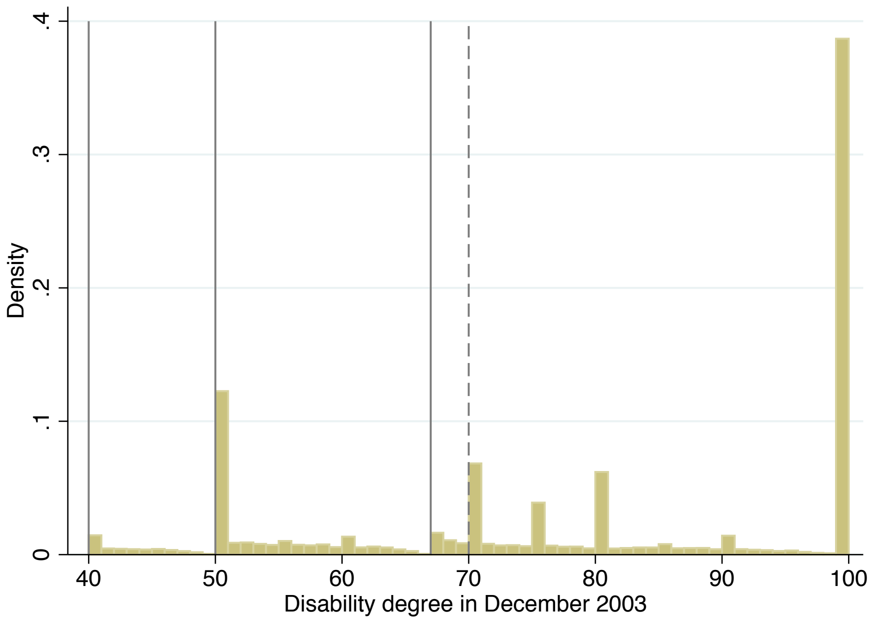
Robust standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Sample differs per column. Outcomes are first differences between the years 2004 and 2003. Earnings are yearly earnings in CHF. Earnings are set to zero if not working. Employment is a dummy variable equal to 1 if the individual has earnings above zero. Disability degree is potential earnings loss as a percentage of potential earnings without disability. Total pension includes main pension, child pension, spousal pension, means tested benefits, and helpless allowances. Controls include dummies for canton of residence and year of observation, marital status, dummy for Swiss citizenship and gender.

Table A.6: Description of never-takers and compliers in the treatment group in 2003

	Never-taker (1)	Complier (2)	Difference (3)
Earnings	5058	5756	-698
	(327)	(586)	(658)
Employment	0.35	0.32	0.03
	(0.02)	(0.03)	(0.03)
Disability degree	67.8	67.8	-0.03
	(0.03)	(0.04)	(0.05)
Total pension CHF	29520	30032	-512
	(373)	(674)	(752)
Income before disability	46938	48856	-1919
	(901)	(1179)	(1695)
Mental illness	0.32	0.24	0.08***
	(0.01)	(0.02)	(0.03)
Bones and organs of movement	0.26	0.39	-0.13***
	(0.01)	(0.03)	(0.03)
Accidents	0.19	0.23	-0.04*
	(0.01)	(0.02)	0.03
Age	46.0	45.6	0.46***
	(0.07)	(0.12)	(0.14)
Married	0.61	0.72	-0.11***
	(0.02)	(0.02)	(0.03)
Foreigner	0.29	0.47	-0.18***
	(0.01)	(0.03)	(0.03)
Female	0.51	0.39	0.12***
	(0.02)	(0.03)	(0.03)
Observations	992	335	1327

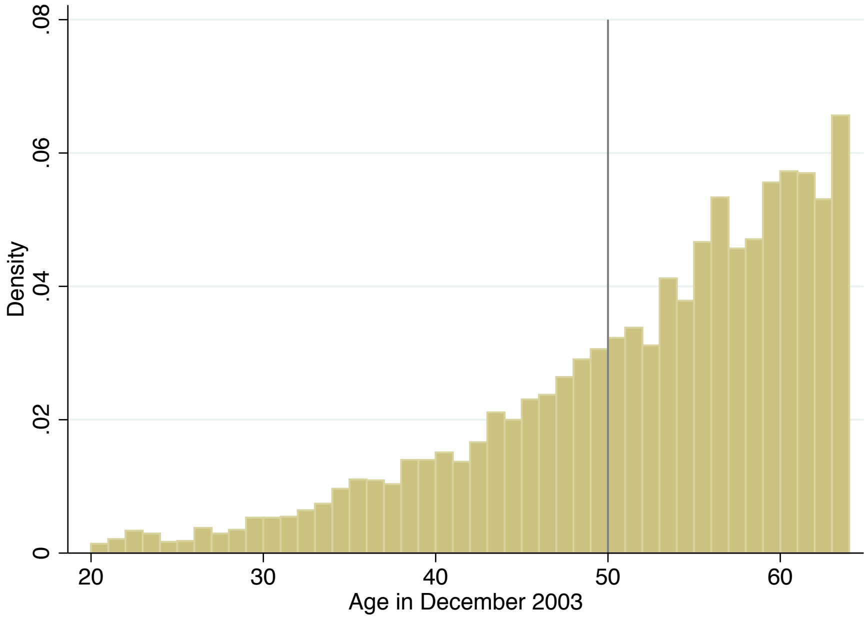
Note: Standard errors in parentheses. Earning is yearly earnings in CHF and equal to zero if the individual is not employed. Employment is a dummy variable equal to 1 if earnings are larger than zero. Disability degree denotes earning loss with disability as a percentage of potential earnings without disability. Total pension includes main pension, child pension, spouse pension, means tested benefits and helpless allowance. In the treatment group with $D = 1$, never-takers are identified having $P_t = 0$ for $t = 2004$. Compliers are identified having $P_t = 1$ for $t = 2004$. This definition assumes away always-takers, who amount to only 3% of the treated DI beneficiaries.

Figure A.1: Histogram for disability degree in December 2003 (individuals aged 42 to 57)



Note: The sample consists of all individuals receiving DI pensions in December 2003 with age between 42 and 57. Disability degree denotes earnings loss as a percentage of potential earnings without disability. Solid vertical lines denote disability degree thresholds to receive a quarter, semi, or full pension, respectively before the reform. The dashed vertical line denotes the disability degree threshold to receive a full pension after the reform in January 2004.

Figure A.2: Histogram for age in December 2003 (individuals with disability degree between 67 and 69)



Note: The sample consists of all individuals receiving DI pensions in December 2003 with a disability degree between 67 and 69. The vertical line in denotes the age threshold above which individuals were not affected by the pension cut of the reform.