Discussion Paper Series

IZA DP No. 18769

July 2026

Measuring the Value of Disability Insurance from Take-Up Decisions

Andreas Haller
Norwegian School of Economics,
CEPR, Cesifo and IZA@LISER

Stefan Staubli
University of Calgary, NBER,
CEPR, IZA@LISER
and RSI Montréal

IZA DP No. 18769 July 2026

Measuring the Value of Disability Insurance from Take-Up Decisions*

Abstract

The central trade-off for designing Disability Insurance (DI) is between providing insurance to those in need while maintaining incentives to work. This paper develops a novel revealed-preference approach to identify the insurance value of DI benefits. We show that comparing the DI take-up response to a change in benefits versus a change in wages identifies the insurance value. Implementing our framework in Canada, we estimate that increasing DI benefits by $1 creates an additional disincentive cost of $0.60 but creates an insurance value of $2.20. Thus, our approach suggests that DI benefits are not overly generous in the Canadian context.

JEL classification

H53, H55, J14, J21

Keywords

disability insurance, take-up, benefits, policy reform

Corresponding author

Stefan Staubli

sstaubli@ucalgary.ca

* We thank Manasi Deshpande, Itzik Fadlon, Giulia Giupponi, Audrey Guo, Lucija Muehlenbachs, Adam Lavecchia, Attila Lindner, Todd Morris, Matthew Notowidigdo, Andreas Peichl, Johannes Spinnewijn, Michael Veall, Alexander Willen, Josef Zweimüller and seminar participants at HEC Montréal, IFAU, McMaster University, University of Geneva, University of Munich, University of Queensland, the Bocconi/CEPR Social Security workshop, the National Tax Association conference, the NBER Social Security Summer Institute, CESifo Area Conference on Public Economics, and the Nordic Public Policy Symposium. We thank Sencer Karademir and Kangyu Qiu for excellent research assistance. This research was supported by funding from the Social Sciences and Humanities Research Council in Canada. The analysis was conducted at the Prairie Regional RDC, as part of the Canadian Research Data Centre Network (CRDCN). This service is provided through the support of the University of Calgary, the province of Alberta, the Canada Foundation for Innovation, the Canadian Institutes of Health Research, the Social Sciences and Humanities Research Council, and Statistics Canada. All views expressed in this work are our own. All remaining errors are our own.

1 Introduction

Disability insurance (DI) programs provide valuable income replacement against losses in work potential due to a disability. At the same time, there are concerns that DI programs distort work incentives and reduce workers' labor supply. While a large body of literature analyzes the impact of DI on labor supply, we know little about the program's value to DI beneficiaries. Yet, without knowledge of both the insurance value and incentive costs, assessing the welfare effects of a change in DI generosity remains elusive.

In this paper, we develop a new method to estimate the value of disability insurance. We show that the insurance value of disability benefits is identified by the ratio of the take-up responses to a change in disability benefits and the take-up responses to a change in wages. The main advantage of our approach compared to existing methods is its broad applicability in settings where data on take-up decisions are available. Moreover, our approach does not require assuming a specific level of risk aversion and can accommodate state-dependent utility functions, unlike the traditional consumption-based approach used to estimate the insurance value.[1]

The intuition why take-up responses identify the insurance value is that the individual's DI take-up decision compares the utility when receiving DI benefits against the utility when working. If a small change in DI benefits leads to a sizable increase in DI take-up, it suggests that higher benefits substantially increase the utility of receiving DI benefits, i.e., the marginal utility of consumption when on DI is high. Conversely, if a minor change in wages triggers a substantial reduction in DI take-up, it suggests that the attractiveness of remaining in the workforce has risen significantly, pointing to a high marginal utility of consumption when working. The relative DI take-up responses to higher benefits and lower wages therefore quantify the ratio of marginal utilities of consumption when on DI and when working, which captures the insurance value of DI benefits. Moreover, the DI take-up response to more generous benefits directly identifies the incentive costs of DI benefits. DI take-up responses are thus sufficient statistics to

[1]The value of a change in social insurance generosity can be calculated as the change in consumption multiplied by workers' risk aversion (Gruber, 1997).

evaluate the incentive-insurance trade-off in the DI program and shed light on whether DI benefits are overly generous.

We implement our method in the context of Canada, allowing us to combine detailed administrative data from 20 percent of the Canadian population with exogenous variations in disability benefits and wages. We start our analysis by studying the DI take-up response to a 1987 reform, increasing disability benefits by 36 percent in the Canadian Pension Plan disability program (CPP-D) to align them with the level of disability benefits in the Quebec Pension Plan disability program (QPP-D). The QPP-D covers residents in the province of Quebec, and the CPP-D covers residents in the rest of Canada. The differential variation in disability benefits in the CPP-D and QPP-D over time enables us to apply a differences-in-differences estimation approach. Our analysis shows that more generous disability benefits lead to more DI take-up. We estimate a disability take-up elasticity of 0.58 (or a $1,000 increase in lifetime DI benefits increases DI take-up by around 0.2 percentage points).

To estimate the impact of wages on disability take-up, we use a Bartik shift-share design that exploits variation in exposure to economic shocks driven by differential industry composition across Canadian census divisions. For robustness, we also exploit variation in labor market conditions across census divisions created by shocks to world oil and gas prices. Census divisions with a high employment share in the oil and gas sector are significantly more affected by oil and gas price shocks compared to census divisions with few workers in the oil and gas sector. Both approaches have been implemented in the literature to estimate DI take-up effects (Black et al., 2002; Autor and Duggan, 2003; Charles et al., 2018; Milligan and Schirle, 2019). We find that adverse economic shocks significantly increase DI take-up in line with the existing literature. In particular, we estimate that a $1,000 reduction in lifetime earnings increases DI take-up by 0.06 to 0.1 percentage points.

Combining the estimated take-up responses to wages and DI benefits, we find that the insurance value of a $1 increase in DI benefits is $2.2 in the Bartik shift-share design. For oil price shocks, the insurance value is even higher, ranging from $2.8 to $3.4. Our estimates thus indicate that the marginal utility gain from an extra dollar is more than double in the disabled state than in the non-disabled state. At the same time, we find that incentive costs from providing more generous disability benefits are sizable but smaller. The fiscal cost of a $1 increase in DI benefits is $1.6, about three-quarters of the insurance value, suggesting that Canadian DI benefits are valuable and not too generous. We also apply our method to the US estimates from Milligan and Schirle (2019) and find that providing one additional dollar in DI benefits comes at a high cost of $2.2. However, the implied insurance value is even higher at $3.4.

There are a couple of nuances to our approach. First, our approach identifies the insurance value of the DI program in the seminal DI model of Diamond and Sheshinski (1995). However, in models with richer heterogeneity and non-additive utility functions, our approach identifies the insurance value of the marginal entrant. We identify conditions under which the marginal entrant's insurance value provides a lower bound for the program's insurance value and present evidence consistent with this interpretation. Second, it is challenging to isolate pure wage effects. The local variation in the exposure to economic shocks we exploit empirically affects wages and employment. In a model extension, we show that exploiting variation in earnings from shocks affecting wages and employment yields again a lower bound for the insurance value. The third nuance relates to the timing of effects. Empirically, we exploit permanent changes in benefits but temporary changes in earnings. Extending our model to a dynamic setting, we show that we can still identify the insurance value from take-up responses by rescaling the DI take-up response of the temporary earnings shock by its impact on the present value of lifetime incomes, which we do in our Canadian implementation.

Our finding that DI benefits are not overly generous is in line with recent papers. They estimate the insurance value of DI using private insurance purchases, consumption drops, and structural models, but the role of take-up decisions for the insurance value has not been recognized. Deshpande and Lockwood (2022) estimate that the value of US disability benefits exceeds a cost-equivalent tax cut by 64%. Cabral and Cullen (2019) infer from purchases of supplemental private insurance in the US that the value of public DI is more than 2.5 times the cost of providing public DI. Using a structural model, Low and Pistaferri (2015) also find that welfare increases with increased DI generosity in the US, implying that the insurance value exceeds the incentive costs of DI benefits.

Meyer and Mok (2019) estimate a consumption drop of 18% for US DI recipients. Using Gruber's (1997) consumption-based formula, this drop implies a relative risk aversion coefficient of at least 6 and 13 to match our insurance value estimates for Canada and the US. The fact that such high levels of risk aversion are required to reconcile consumption-based estimates with those from revealed preference approaches is consistent with findings in the UI literature, which compare insurance values across methods in the same context (Landais and Spinnewijn, 2021).[2] In contrast to the U.S. evidence of high insurance values, Seibold et al. (2025) find that willingness-to-pay for DI is relatively low based on the limited take-up of private insurance in the German context.

The extensive literature on the determinants of DI take-up has interpreted strong take-up responses to either a change in benefits or a change in wages as a sign of moral hazard limiting the effectiveness of the DI program. We offer a new interpretation and show that the ratio of these two take-up margins is critical for the welfare implications. The branch of this literature that focuses on DI benefit generosity has sought to estimate the impact of benefit levels on labor supply (for a review of the literature, see Bound and Burkhauser, 1999; Low and Pistaferri, 2020). Most closely related to our study is Gruber (2000) who exploits the same DI benefit reform in Canada. He finds that a 36% increase in DI benefits induced an 11.5% increase in non-employment among 45-59-year-old men in the first two years after the reform. Instead, our paper estimates the impact of this reform on DI take-up, which is vital for inferring the insurance value of DI benefits. Moreover, we estimate the impacts on the entire working-age population up to 13 years after the reform.

Our approach to estimating the impact of wage changes on DI take-up builds on several well-known US studies. Of particular relevance is the study by Autor and Dug-

[2] Landais and Spinnewijn (2021) require a coefficient of relative risk aversion between 4 and 8 to reconcile the high insurance values implied by their marginal propensity to consume and revealed preference estimates with relatively modest consumption drops.

gan (2003), which uses a similar industry shift-share design to study the impact of labor market shocks across US states on DI take-up. Milligan and Schirle (2019) apply a shift-share design in both the Canadian and US contexts. Another set of US studies estimates the impact of wage changes on DI take-up using booms and busts in coal mining (Black et al., 2002) and oil and gas production (Charles et al., 2018). Since Canada is a large oil and gas producer, we use oil and gas price shocks as a robustness check and find quantitatively similar estimates as Charles et al. (2018).

Methodologically, our paper contributes to the sufficient statistics literature that identifies the insurance value of social insurance benefits from observed behavior. Recent studies have developed methods to estimate the insurance value of unemployment insurance (UI) (Landais and Spinnewijn, 2021; Hendren, 2017; Landais, 2015; Chetty, 2008; Shimer and Werning, 2007; Chetty, 2006; Gruber, 1997). Our approach is conceptually related to Chetty (2008), who shows that the insurance value of UI benefits can be identified as the ratio of job search effort responses to changes in unconditional cash transfers and changes in UI benefits. Chetty (2008) emphasizes the exit margin (job search responses) to identify the value of UI benefits. Similarly, Shimer and Werning (2007) show that the response of reservation wages, which make an individual indifferent between working and remaining unemployed and govern exit from unemployment, is sufficient to characterize optimal benefit generosity. We show that take-up responses reveal the insurance value of social insurance programs in which the entry margin is important and individuals choose between working and receiving benefits, such as in DI, sick leave, or old-age pensions. Our approach can therefore be applied to other social insurance programs and requires only data on take-up, benefits, and wages, which are widely available in many countries.

Similar in spirit to our analysis, Eisenhauer et al. (2015) exploit program participation decisions, together with instruments that shift program entry, to identify the benefits, costs, and net surplus of program participation. While their approach focuses on the value of the program, ours identifies the value of increasing program generosity.

The paper is organized as follows. The next section presents a model of disability insurance and derives formulas for optimal disability benefits. Section 3 describes the data and the institutional background. Sections 4 and 5 present the empirical results on changes in DI benefits and wages. Section 6 estimates the insurance value and the incentive costs of more generous DI benefits. Section 7 concludes.

2 Model: The Value of DI Benefits

To illustrate that take-up decisions identify the insurance value of DI benefits, we start with a simple version of the seminal DI model by Diamond and Sheshinski (1995). In the simple model, individuals work or receive DI benefits and only differ in their disability levels, which enters utility additively. In this model, the relative take-up responses of higher DI benefits vs. lower wages identify the insurance value of DI benefits.

We then extend the simple framework in three ways. We add a third labor market state (other benefits) and endogenous job search, we allow for non-additively separable utility functions and richer heterogeneity beyond disability levels, and we provide a dynamic version of the model.

Setup. Consider a continuum of agents living for one period. Agents differ only in their level of disability θ, modeled as a random draw from a continuous distribution $F(\theta)$. An agent can decide to work or to apply for DI. An agent with disability level θ enjoys utility $u(w-\tau)-\theta$ if she works and utility $v(b)$ if she receives DI benefits. The outcome of a DI application is uncertain; an application is accepted with probability $p(\theta)$, where $p'(\theta) > 0$. Assuming no application costs, an agent applies if she prefers receiving DI benefits over working (i.e. if her disability level θ is high).[3] The "marginal entrant", the agent who is indifferent between applying for DI benefits and remaining employed, has disability

$$\theta^A = u(w-\tau) - v(b). \tag{1}$$

[3]The assumption of no application cost is not critical for our result. With application costs ψ, the marginal entrant is determined by $p(\theta^A)\left(v(b) - \left[u(w-\tau)-\theta^A\right]\right) - \psi = 0$ and we can use the same approach to measure the insurance value of DI.

Agents with disability $\theta \geq \theta^A$ apply for DI, while agents with disability $\theta < \theta^A$ remain employed.[4]

Welfare. The government maximizes a utilitarian welfare function subject to a balanced budget constraint:

$$\max_{b,\tau} V(b,\tau) \text{ s.t. } G(b,\tau) = 0, \tag{2}$$

where $V(b,\tau)$ denotes the aggregate indirect utility function for a given DI benefit level b and tax τ,

$$V(b,\tau) = \int_0^{\theta^A} [u(w-\tau)-\theta]\, dF(\theta) + \int_{\theta^A}^{\infty} [p(\theta)v(b) + (1-p(\theta))(u(w-\tau)-\theta)]\, dF(\theta). \tag{3}$$

The first term of equation (3) measures the utility of individuals who do not apply to DI and work. The second term sums up the utilities of accepted DI applicants, $p(\theta)v(b)$, and of rejected DI applicants, $(1-p(\theta))(u(w-\tau)-\theta)$. $G(b,\tau)$ denotes total fiscal revenue, which can be written as

$$G(b,\tau) = (1-DI)\cdot\tau - DI\cdot b, \tag{4}$$

where

$$DI \equiv \int_{\theta^A}^{\infty} p(\theta)dF(\theta) \tag{5}$$

denotes the share of individuals on DI benefits.

[4] The original framework of Diamond and Sheshinski (1995) includes a third labor market state, "other welfare benefits." For notational simplicity, we abstract from this state in the baseline model. Our approach is based on the behavior of marginal entrants, whose decision is characterized by equation (1) as in Diamond and Sheshinski (1995), and therefore applies irrespective of the presence of other welfare benefits, as long as the relevant outside option for the marginal entrant is work. In Appendix A.3 and at the end of this section, we discuss the implications when other welfare benefits become the relevant outside option for some marginal entrants.

Optimal DI Benefits. The optimal level of DI benefit trades off the more generous insurance against the higher program costs. The solution to problem (2) fulfills the classic Baily-Chetty formula

$$\underbrace{\frac{v'(b)}{u'(w-\tau)}}_{\text{Insurance Value}} = \underbrace{1 + \frac{\varepsilon_{DI,b}}{1-DI}}_{\text{Multiplier}}, \tag{6}$$

where $\varepsilon_{DI,b} \equiv \frac{dDI}{db}\frac{b}{DI}$ is the elasticity of DI take-up with respect to DI benefits.[5] The LHS of (6) measures the insurance value of DI benefits—the relative valuation of \$1 in the working state versus \$1 when on disability benefits. The RHS measures the total cost of a \$1 transfer between the employment and disability benefit state.

Inflow Responses Reveal Insurance Value of DI. We now explore how the insurance value can be identified from the relative DI take-up responses to higher DI benefits versus lower wages. The DI take-up effect of higher DI benefits is

$$\frac{\partial DI}{\partial b} = f(\theta^A) \cdot p(\theta^A) \cdot [v'(b)], \tag{7}$$

which follows from definition (5) and the application decision in (1). Similarly, the DI take-up effect of a reduction in wages is

$$-\frac{\partial DI}{\partial w} = f(\theta^A) \cdot p(\theta^A) \cdot [u'(w-\tau)]. \tag{8}$$

Together, the relative DI take-up response to benefits and wages measures the insurance value:

$$-\frac{\partial DI}{\partial b} / \frac{\partial DI}{\partial w} = \frac{v'(b)}{u'(w-\tau)}. \tag{9}$$

Measuring the insurance value from take-up decisions is a new idea. The main advantage of this approach is that DI take-up effects can be estimated in many settings with standard administrative data. Two influential strands of literature estimate the take-up effects of changes in DI benefit generosity and the DI take-up effects of labor market

[5]The formal derivation of the optimality condition (6) is presented in Appendix A.1.

shocks (reduced wages). Our method combines these two strands of the empirical DI literature. Through the lens of our theory, we provide a new interpretation of the relative take-up effects as a measure of the insurance value of DI benefits.

Intuition. Individuals apply to DI benefits when the utility they derive from being on DI surpasses the utility when working. When a slight adjustment in DI benefits leads to a sizable change in DI take-up, it indicates that a small benefit increase substantially boosts the utility derived from receiving DI benefits, i.e., the marginal utility of consumption when on DI is high. Conversely, if a minor wage increase significantly lowers DI take-up, it suggests that remaining in the workforce has become much more attractive, pointing to a high marginal utility of consumption when working. The relative DI take-up responses to higher benefits and lower wages therefore quantify the ratio of marginal utilities of consumption when on DI and when working (insurance value).

Another way to view this result is that DI benefits and wages are state-contingent transfers. One only receives DI benefits while on the DI program and wages only when working. If workers can perfectly self-insure against disability risk, they should be indifferent whether DI benefits or wages increase. The take-up responses to changes in benefits and wages are then the same because workers can transfer income across the states freely; the marginal utilities of consumption when on DI and when working are the same. In case of imperfect insurance the relative DI take-up responses identify the wedge in the ability to insure against disability risk.

How General Is This Result? A crucial question is whether our approach to identify the insurance value is limited to the specific model above or applies more generally. The model above is too simple in three main respects. First, the baseline model abstracts from several important dimensions of heterogeneity, like variation in wages, DI benefits, and unearned income, and it assumes additive separability between consumption and disability. Appendix A.2 shows that, in richer environments allowing for preference non-separability and additional heterogeneity, our approach identifies the insurance value of marginal entrants. The key question then is how representative the marginal entrant's insurance value is for the average DI recipient's insurance value.

A natural conjecture is that the marginal entrant provides a lower bound. In Appendix A.2, we characterize the conditions under which this holds. If marginal utility depends only on consumption and utility is concave, the lower-bound result follows when marginal entrants have weakly higher consumption than average DI recipients when receiving DI benefits and weakly lower consumption than average taxpayers when working. While we do not observe consumption directly, our complier analysis in Section 5 shows that marginal entrants are disproportionately drawn from the lower end of the income distribution and are similar to average DI recipients in pre-DI income and mortality. This suggests that they are comparable to average DI recipients when on DI but worse off than the average taxpayer when working, which is consistent with the lower-bound interpretation.

When marginal utility also depends on disability severity, the result is more nuanced. If marginal utility increases with disability severity, the lower-bound result is reinforced, as more severely disabled individuals both value consumption more and are more likely to receive DI.[6] By contrast, if marginal utility decreases with severity, the lower-bound result need not hold.

Overall, our complier analysis and standard assumptions in the DI literature regarding separability are consistent with interpreting the marginal entrant's insurance value as a lower bound. However, whether this holds in practice is ultimately an empirical question. In settings with consumption data, this could be examined in future work by comparing consumption drops between marginal and average entrants.

A second limitation of the baseline model is that it features only two labor market states (working or receiving DI benefits). This is not a theoretical restriction: as long as the relevant outside option for the marginal type is work, equation (9) and the results from the baseline model continue to apply. Appendix A.3 presents two extensions that

[6] Low and Pistaferri (2015) also assume in their structural model that marginal utility of consumption is higher when individuals face a work limitation. However, the empirical evidence on how marginal utility varies with health is inconclusive. Finkelstein et al. (2013) find that marginal utility declines in worse health states, whereas Lillard and Weiss (1997) find the opposite. Brown et al. (2016) find that, on average, working-age individuals value healthy and unhealthy states similarly when facing a potential disability that limits labor market participation, but document substantial heterogeneity in the degree of state dependence across individuals.

introduce a third labor market state. The first allows for the possibility that economic shocks affect both wages and employment, and the second allows for benefit substitution.

The first extension incorporates unemployment benefits and endogenous job search to address an empirical challenge. We exploit local labor market shocks to identify pure wage effects, but such shocks may also affect separation and job-finding rates. The extended model shows that when shocks affect both wages and employment, the DI take-up response should be rescaled by the shocks' impact on total income, as income captures the combined effect of changes in wages and employment. This rescaling yields a lower bound for the marginal entrant's insurance value when the monetized utility loss from job loss exceeds the corresponding income loss—a condition that holds under standard assumptions such as concave, state-independent utility and an unemployment insurance replacement rate below 100 percent.

The second extension addresses benefit substitution: changes in DI benefit generosity may induce inflows from both work and other welfare programs. To capture inflows from both states, we allow individuals to differ in disability severity and DI application costs. Individuals with high application costs may not apply for DI even when severely disabled and instead rely on other welfare programs. As a result, benefit changes can generate inflows both from individuals whose outside option is work and from those whose outside option is other benefits.

This extension shows that benefit substitution creates an upward bias in our insurance value estimate, as benefit changes generate additional inflows from other welfare programs while wage changes do not. The magnitude of this bias can be quantified empirically as the ratio of reform-induced DI inflows from other welfare programs to the DI inflow response to a wage change. Although we do not directly observe receipt of other welfare programs, Table 1 shows that most of the estimated DI inflow originates from employment rather than from non-employment, suggesting that benefit substitution is limited in our setting.

The third shortcoming of the simple model is that it is static. Appendix A.4 shows that our approach generalizes to a dynamic setting. The main challenge is linking the model to the empirical variation: the change in DI benefits is permanent, whereas the wage shocks we exploit are temporary. We show that the insurance value can be identified from permanent benefit changes and temporary wage changes by rescaling the DI take-up response by the temporary wage shock's impact on the present value of lifetime earnings.

This rescaling requires individuals to have correct beliefs about the persistence of the income shocks. If individuals overestimate persistence, our estimated DI take-up response to income shocks is too large, leading to a downward bias in the estimated insurance value.[7]

A growing literature shows that individuals systematically misperceive the persistence of income shocks. Evidence on extrapolation bias and diagnostic expectations suggests that individuals tend to overreact to recent income changes, treating transitory shocks as too persistent (e.g., Bordalo et al., 2022; Rozsypal and Schlafmann, 2023; D'Acunto et al., 2024). In light of this evidence, our estimates can be interpreted as a lower bound: if individuals overreact to temporary income shocks, the observed take-up response overstates the response to the true lifetime-income change, leading us to understate the insurance value.

3 Institutional Background and Data

Our setting to estimate the value of disability insurance benefits is Canada. This section describes the Canadian public long-term DI program, the variation in DI benefits and earnings we exploit in the empirical analysis, and the data.

[7]To illustrate this point, suppose individuals believe that a shock changes their lifetime income by $d\tilde{w}$, while the true change is dw. Empirically, we observe the DI take-up response with respect to the perceived income shock, but rescale it by the true shock size; that is, we estimate $\frac{dDI(\tilde{w})}{dw}$. However, the revealed preference approach requires $\frac{dDI(\tilde{w})}{d\tilde{w}}$. If individuals overestimate the persistence of income shocks ($d\tilde{w} > dw$), we overstate the DI take-up response with respect to income shocks, since $\frac{dDI(\tilde{w})}{dw} > \frac{dDI(\tilde{w})}{d\tilde{w}}$. The denominator in our insurance-value estimator is therefore too large, causing us to underestimate the insurance value. The opposite logic applies if individuals underestimate the persistence of income shocks.

3.1 Institutional Background and Policy Variation

A peculiarity of Canada is the existence of two separate DI programs, both established in 1966 and financed by a payroll tax. Quebec, the second largest province with about a fourth of the Canadian population, desired to retain control over the pension plan's design and created its own program, the Quebec Pension Plan Disability (QPP-D). In contrast, the other provinces are all covered by the Canadian Pension Plan Disability (CPP-D) program.[8] Today, the programs are nearly identical concerning eligibility criteria and level of benefits, but historically QPP-D benefits were higher until a 1987 reform increased the CPP-D benefits. We exploit this reform to estimate the impact of benefit generosity on DI take-up.

To qualify for DI benefits, an individual needs to suffer from a prolonged—lasting at least one year—mental or physical disability that prevents pursuing any substantially gainful employment. Eligibility is determined by a medical examiner and individuals who are denied can appeal their decision. In 1989, the overall award rate was 68 percent under the CPP-D and 70 percent under the QPP-D (Gruber, 2000). Eligibility also depends on the contribution time since age 18. Before 1987, individuals had to contribute 5 out of the last 10 years or one-third of the contribution period. The CPP-D relaxed contribution requirements in 1987, allowing workers who contributed 2 of the past 3 years to also qualify. It tightened the requirements in 1997 (4 contribution years in the last 6 years) and relaxed them again in 2005 (3 contribution years in the last 6 years or 25 years in total). The QPP-D changed the contribution period only once. Since 1993 individuals are eligible if they contributed half of the contribution period, 5 of the past 10 years, or 2 of the last 3 years. Once benefits are awarded, DI recipients receive monthly payments until they die, return to work, or reach the retirement age.

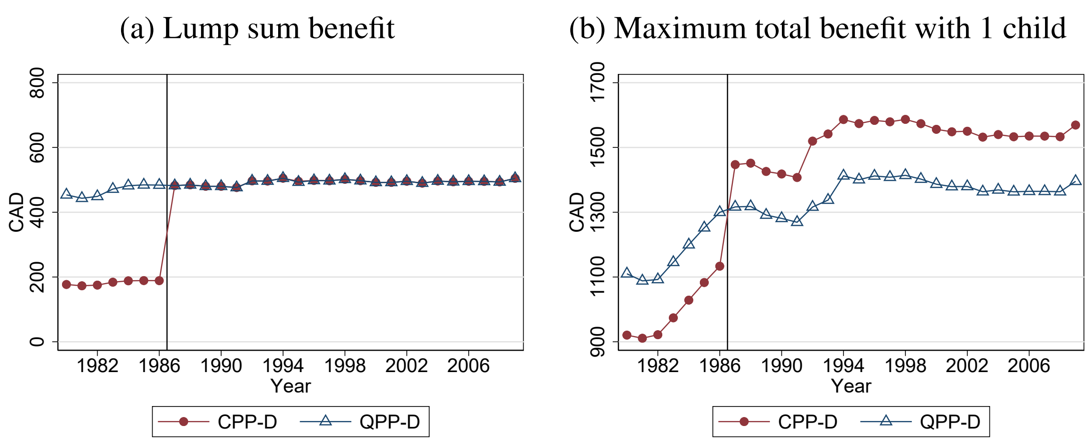
Variation in Potential DI Benefits. DI benefits in the CPP-D and the QPP-D consist of three parts: a lump-sum benefit identical for all eligible recipients, an earnings-related benefit, and a child allowance, which is a fixed amount per child under age 18. The earnings-related benefit is calculated in the same way in both programs, but after

[8]For detailed descriptions of the CPP-D and the QPP-D, see Torjman (2002), Baker and Milligan (2012), and Campolieti and Riddell (2012).

1972 the lump-sum component grew much faster in the QPP-D compared to the CPP-D.[9] By 1986, the lump-sum transfer in the QPP-D was almost three times as large as in the CPP-D, as illustrated in Panel (a) of Figure 1.

In an effort to align the two programs, in January 1987 the government raised the CPP-D lump-sum benefit to the same level as in the QPP-D. This change increased the annual CPP-D benefits by about CAD 3,500 (in 2019 dollars) or about 36 percent. Panel (b) of Figure 1 shows that before the reform the maximum CPP-D benefits were lower than in the QPP-D, but the 1987 reform raised CPP-D benefits above the level in the QPP-D. After the reform, the average monthly DI benefit payments moved in parallel, except for minor differences between 1992 and 1994. These differences arose because the CPP-D increased the child component in 1992, while the QPP-D followed only in 1994.

Figure 1: Maximum Monthly DI Benefit Payments in CPPD and QPPD over Time



Notes: The figure shows the maximum lump sum DI benefit (panel a) and maximum total DI benefit with one child (panel b) in the CPP-D and QPP-D (in 2019 Canadian Dollars). Before the reform in 1987 DI benefits in CPP-D were less generous than in QPP-D. The adjustment in the lump-sum component of the CPP-D benefits in the 1987 reform raised CPP-D benefits above the QPP-D level. Pre- and post-reform the CPP-D and QPP-D benefits evolve in parallel. Numbers are based on the CPP STATS BOOK 2019 and "Évolution de la clientèle de la rente d'invalidité de 1970 à 2010" (Diarra et al., 2015).

The 1987 reform implemented two additional changes to the CPP that could matter for our analysis. As discussed above, the reform relaxed the contribution requirement, permitting workers with 2 contribution years in the last 3 years to qualify for bene-

[9]For the earnings-related benefit the DI recipients earnings history is inflated by a wage index and the lowest 15 percent of monthly real earnings are dropped. The earnings-related benefit is then calculated as 18.75 percent of the average monthly earnings of the remaining earnings history.

fits. To isolate the benefit generosity effect, we would ideally restrict the sample to individuals who fulfill pre- and post-reform eligibility criteria. But since our data start only in 1982, we cannot verify the pre-reform contribution requirement over the last 10 years. Instead, we look at individuals in 1992 for whom we can observe the contribution history over the last 10 years. We find that only 2.3% would qualify with 2 out of 3 contribution years alone. Almost 87% would already qualify under the pre-reform contribution requirement, while 11% never qualify. Hence, the potential impact of the relaxation in contribution requirements is small. The 1987 reform also lowered the early retirement age in the CPP from age 65 to 60. We therefore restrict our sample to ages 15 to 59. Still, forward-looking individuals might adjust labor supply before age 60, but such anticipatory effects are likely small, as previous literature finds that the change had little impact on the labor supply of 60-64-year-olds (Baker and Benjamin, 1999; Staubli and Zhao, 2023). Moreover, the QPP lowered its early retirement age already in 1984, and we find that this change had no impact on labor supply or DI take-up before age 60.

Variation in Potential Earnings. To estimate the impact of potential earnings on DI take-up, we combine granular variation in industrial composition across Census Divisions (CDs) with national-level changes in employment to predict local shocks in potential earnings. Census Divisions are established by provincial law and are roughly similar to US counties. Canada has about 260 CDs, which differ significantly in their industry composition. For example, manufacturing industries such as auto-making, food and beverage, and fabricated metals are concentrated in CDs in Ontario, while the technology sector is strong in CDs in Quebec. The variation implies that industry-specific shocks will impact some CDs much more than others. For example, a demand shock for cars affects CDs in Ontario but have little impact on CDs in Quebec.

We leverage the cross-CD variation in industrial composition and national-level changes in employment using a Bartik shift and share instrumental variable approach. Our main specification exploits variation in industry composition across three-digit industries (102 industries) similar to Autor and Duggan (2003). We probe the robustness of our results using variation in oil and gas employment across CDs as in Charles et al.

(2018), paired with national changes in oil and gas employment or prices. We use both national employment and prices changes, since Canadian oil prices were under price controls from 1974 to 1985.

Oil and gas employment is concentrated in CDs in the provinces of Alberta, Saskatchewan, British Columbia, and Newfoundland and Labrador. Appendix Figure C.1 displays the real oil price trends between 1982 and 2016.

3.2 Data

We use individual tax return data from the Longitudinal Administrative Databank (LAD), a representative panel of 20 percent of Canadian tax filers between 1982 and 2019. Individuals who are selected into the LAD are followed each year they file a return and they can be linked across years using an anonymous identifier. The LAD contains information on earnings, government transfers—including Canada and Quebec pension plan benefits—taxes, and demographics. The data also contain detailed geographic information including the census division and province of residence, which we use to infer whether an individual is covered by the CPP-D or QPP-D.[10]

We measure DI benefit receipt by whether an individual receives a public pension. The data does not distinguish between the type of public pension benefits received, but the only available benefit before age 60 other than DI are survivor benefits. The 1987 reform did not change the generosity of survivor benefits. We would thus not expect any change in the take-up of survivor benefits around the reform. Since the data record the spouse's year of death, we can approximate the receipt of survivor benefits using spousal deaths. In a robustness check, we set DI benefit receipt to zero for individuals who start receiving benefits within a year their spouse died. The results are nearly identical.

One drawback of the LAD is that it does not record an individual's occupation or industry, which we need to calculate the CD industry shares. To address this issue, we

[10]The geographic information in each year is based on the latest available Census for each given year. To avoid measurement error from changes in Census Divisions, we construct time-consistent definitions of Census Divisions that line up with the 1981 Census definitions.

pool data from the long-form Canadian census for the years 1981, 1986, 1991, 1996, 2001, 2006, 2011, and 2016. The long-form census contains the 3-digit industry of employment for 20 percent of the population.

Analysis Samples. To study the effect of more generous DI benefits, we focus on 15-59-year-old individuals who continuously file taxes from 1982 until 1992 or until they die, whichever is first.[11] We impose this sample restriction to have a balanced sample for the main period of analysis.

To study the effect of changes in potential earnings, we focus on 15-59-year-old individuals in the LAD from 1982 to 2016. We observe individuals in five-year intervals because the Canadian Census is only conducted every 5 years and we need the Census to calculate industry shares.[12]

4 Impact of Benefit Generosity

4.1 Estimation Strategy

The 1987 reform increased CPP-D benefits while leaving QPP-D benefits unchanged. We exploit the policy-induced variation in benefits in a difference-in-differences (DiD) design, comparing the change in an outcome variable in the Rest of Canada (RoC) with the change in the same outcome variable in Quebec over time. This comparison can be implemented with the following regression:

$$Y_{ipt} = \alpha + \sum_{s=1982, s\neq 1986}^{2000} \beta_s (I[p = RoC] \cdot I[t = s]) + \theta_p + \pi_t + X'_{ipt}\delta + \epsilon_{ipt}, \qquad (10)$$

where Y_{ipt} is an outcome variable such as an indicator for DI take-up of individual i living in province p in year t, $I[p = RoC]$ is a dummy equal to 1 if an individual

[11] Appendix Table B.1 reports summary statistics for this sample, separately for Quebec and the Rest of Canada before and after 1987.

[12] Since the LAD only starts in 1982, we combine Census industry shares in 1981 with LAD variables in 1982. Appendix Table C.1 provides summary statistics for this sample.

lives in the RoC, $I[t = s]$ is an indicator for the observation being in year s, θ_p are province fixed effects, π_t are year fixed effects, and X_{ipt} is a vector of demographic and labor market characteristics (e.g., age, gender, and the provincial unemployment rate). Gruber (2000) studies the effect of the 1987 reform on labor force non-participation for 45-59 year old men using the Canadian Survey of Consumer Finances from 1985-1989 and a "2x2" difference-in-differences design. We build on his work by considering the impact on DI take-up—a central sufficient statistic for welfare analysis—for the full labor force population of 15-59 year old men and women and we estimate the dynamic effects in each year between 1982 and 2000.

The coefficients of interest are the β_s, each β_s-coefficient measures the average causal effect of the reform-induced benefit increase in year s relative to the base year, 1986. The pre-reform β_s-coefficients ($s < 1987$) provide pre-tests for spurious trends. They should not be statistically significant if the identification assumption of parallel trends holds, although they could pick up anticipation effects. Such effects are unlikely, as the reform was only enacted six months before it became effective. We cluster the standard errors at the census division level (roughly 260 clusters).

For our main effects of the reform, we use the 1991-coefficient (β_{1991}), which measures the causal impact of higher DI benefits on individuals in the RoC relative to those in Quebec in 1991. We stop in 1991 for two reasons. First, the 1991-coefficient measures the impact five years after the reform, which matches the time window of the wage effects analysis that also looks at five-year changes due to the census's five-year intervals. Second, the CPP-D lifted the time limit on late applications in 1992. Initially, individuals who applied for CPP-D benefits more than 15 months after the onset of a disability were automatically denied. The 1992 change abolished the 15-month time limit.[13] Moreover, the CPP-D increased the child component of DI benefits in 1992. The QPP-D followed only in 1994, leading to slight differences in benefit generosity trends over these two years. Still, we also assess the long-run impacts of more generous benefits by documenting the effects until 2000 but are more cautious in interpreting the estimates after 1991 as causal impacts of the 1987 reform.

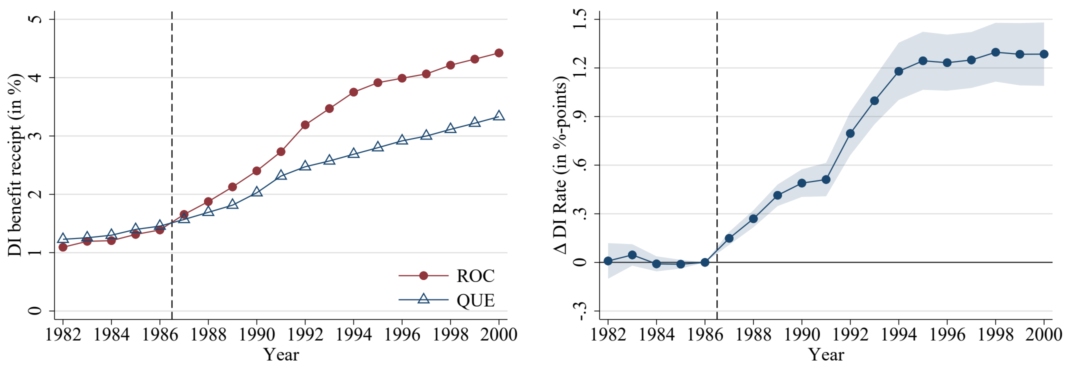
[13] See table 1 in Campolieti and Riddell (2012) for an excellent overview of the policy variation after 1992.

4.2 Empirical Results

We start our analysis by plotting the β_s-coefficients from regression (10). Figure 2(a) shows the trends in DI take-up in RoC (red circles) and Quebec (blue diamonds) from 1982 to 2000, while Figure 2(b) shows the difference in DI take-up between RoC and Quebec from equation (10). The figures show that higher DI benefits induce more entry into DI. Before the reform, DI take-up is similar between RoC and Quebec, providing evidence for parallel trends in DI take-up. When the reform becomes effective in 1987, DI take-up increases in RoC relative to Quebec. In 1991, DI take-up is about 0.5 percentage points higher in RoC relative to Quebec. The effect starts to level off in 1994 with a gap of about 1.2 percentage points in DI take-up rates. As a robustness test, we set DI benefits receipt to zero if the spouse dies within a year of benefit take-up to account for the receipt of survivor benefits. Appendix Figure B.1 shows that the patterns and magnitudes of the estimates are almost identical.

Figure 2: Trends in DI Take-Up Between RoC and Quebec by Calendar Year

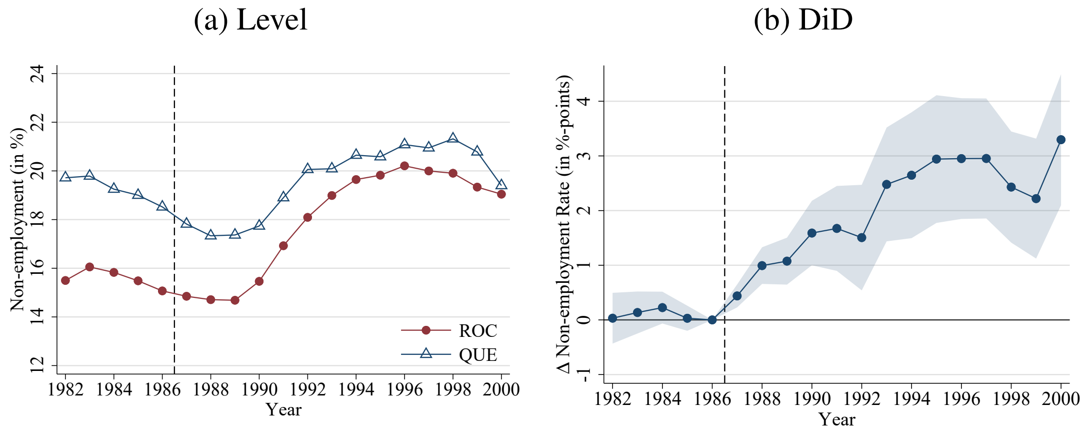
(a) Level (b) DiD



Notes: The figure plots the average DI take-up rate in panel (a) and the average difference in DI take-up obtained from the econometric specification in (10) in panel (b). The shaded area represents the 95-percent confidence interval.

Figure 3 plots trends and the difference in non-employment—an indicator for zero earnings on the tax return—between RoC and Quebec. Not being able to pursue substantial gainful employment is an essential eligibility requirement for DI benefits. The figures show that the reform reduced employment: the rate of non-employment rises sharply in RoC compared to Quebec after the reform. The non-employment effect is almost twice as large as the DI take-up effect (but also less precisely estimated).

Figure 3: Trends in Non-Employment Between RoC and Quebec by Year



Notes: The figure plots the average non-employment rate in panel (a) and the average difference in non-employment obtained from the econometric specification in (10) in panel (b). The shaded area represents the 95-percent confidence interval.

Table 1 presents the reform effect five years after the reform (β_{1991} from regression 10). More generous DI benefits raise DI benefit receipt by 0.527 percentage points. The estimate for DI benefit receipt implies an elasticity of DI benefit receipt with respect to DI benefits of 0.58. This elasticity identifies the fiscal externality in our model for optimal DI benefits, similar to the standard Baily-Chetty model for unemployment insurance. It implies that a one-dollar increase in DI benefits costs taxpayers a total of 1.6 dollars because of induced entry into DI.

Table 1 also decomposes the overall DI inflow effect into transitions from employment and non-employment, where inflows from non-employment are defined as entries from individuals with no labor income in the previous two years.[14] This measure is intended to capture individuals who are not active in the labor market and may rely on other welfare programs. The decomposition shows that most of the DI inflow effect is driven by individuals who were active in the labor market in the previous two years (87 percent of total inflows), with only a small share coming from longer-term non-

[14] Appendix Figure B.2 presents the corresponding event-study by prior employment status. It shows that only a small share of DI inflows initially originates from non-employment, although this share increases over time, particularly after 1992, which lies outside our main period for the welfare analysis.

Table 1: Impacts of 1987 Reform

	DI received	DI from employment	DI from non-employment	Non-employment
Coeff. estimate (β_{1991})	0.527***	0.458***	0.070***	1.689**
	(0.061)	(0.049)	(0.020)	(0.677)
Mean ($\overline{y}$)	1.84	1.36	0.48	14.21
Elasticity $\left(\beta_{1991} \cdot \frac{\overline{b}}{\overline{y}}\right)$	0.580***	0.682***	0.293***	0.241**
	(0.067)	(0.073)	(0.084)	(0.096)
No. Observations	17,543,255	17,543,255	17,543,255	17,543,255

Notes: The table reports the reform effect in the fifth year after the reform for different outcomes (the β_{1991}-coefficient from equation 10). "DI received" is an indicator equal to one if an individual receives a public disability pension. "DI from Employment" is an indicator equal to one if the individual receives a public disability pension and had positive labor income in the two years before entering DI. "DI from Non-Employment" is an indicator equal to one if the individual receives a public disability pension and had no labor income in the two years before entering DI. "Non-employment" is an indicator equal to one if the individual has not labor income. Coefficient estimates and standard errors are multiplied by 100 and should be interpreted as percentage points. Standard errors are reported in parentheses and are clustered at the Census Division. The sample includes 15-59 year-old individuals from 1982 to 1991. All specifications include year fixed effects, province fixed effects, age and its square, and the provincial unemployment rate. Levels of significance: $^{\star}10\%$, $^{\star\star}5\%$, and $^{\star\star\star}1\%$.

employment. This pattern suggests that benefit substitution is limited in our setting.[15]

Non-employment increases by 1.69 percentage points. The implied elasticity of non-employment with respect to disability benefits is 0.24, which is consistent with Gruber (2000). He finds non-participation elasticities of 0.28–0.36 among older male Canadian workers.

5 Impact of Wage Shocks

5.1 Estimation Strategy

To assess the effects of a change in earnings on disability take-up, we follow previous literature (Black et al. (2002); Autor and Duggan (2003); Charles et al. (2018)) and estimate first-difference regressions of the form:

$$\Delta y_{ict} = \alpha + \beta(\Delta E_{ict}) + \lambda_t + \delta \Delta X_{ict} + \varepsilon_{ict}, \tag{11}$$

[15] In our data, we do not observe take-up of other welfare benefits such as provincial social assistance before 1992. We therefore approximate benefit substitution by the DI inflow coming from individuals out of the labor force in the last two years.

where Δ denotes the first difference operator over five-year intervals. The primary outcome variable Δy_{ict} is the change in DI claiming for individual i living in census division c in year t, ΔE_{ict} is the change in annual earnings in \$1,000, λ_t is a vector of year dummy variables, ΔX_{ict} is a vector of changes in control variables (age and age-squared), and ε_{ict} are any unobserved factors affecting DI claiming such as tastes for work.[16] We estimate the regression in levels because our sufficient-statistics formula requires the DI take-up response to a dollar change in earnings. The main parameter of interest, β, thus measures the effect of a \$1,000 change in potential earnings on DI claiming, which is the earnings-response parameter required by the model.

We expect that $\beta < 0$ as higher potential earnings should reduce the likelihood of claiming DI benefits. A concern in estimating β using equation (11) is that ΔE_{ict} is endogenous as changes in earnings are likely correlated with changes in unobserved factors affecting DI claiming. For example, the OLS estimate of β will be biased downward if unobserved tastes for work, ε_{ict}, are positively correlated with earnings and negatively correlated with disability claiming. We require plausibly exogenous changes to potential earnings to surmount this endogeneity concern. We follow Autor and Duggan (2003) and instrument for ΔE_{ict} using a Bartik shift-share design that exploits variation in industry composition across census division (CD) and national-level changes in employment to predict CD-specific employment growth. Following the canonical example in Goldsmith-Pinkham et al. (2020), we construct the shift-share measure using three-digit industries:

$$\Delta \hat{\gamma}_{ct} = \sum \omega_{kct-5} \cdot \Delta \gamma_{\not{k}t}, \tag{12}$$

where $\Delta \hat{\gamma}_{ct}$ is the predicted log employment change for each CD c between $t-5$ and t, ω_{kct-5} is the share of employment in industry k in CD c in year $t-5$, and $\Delta \gamma_{\not{k}t}$ is the log change in the national employment share of the three-digit industry k. As in Autor and Duggan (2003), we exclude each CD's industry k employment when calculating $\Delta \gamma_{\not{k}t}$, hence the subscript $\not{k}$.

To probe the robustness of our results, we also exploit variation in oil and gas em-

[16] We measure annual earnings as the tax-reported annual labor earnings (including self-employment income), deflated using the national CPI (base year 2019).

ployment across CDs, paired with national-level changes in oil and gas employment and prices, similar to Charles et al. (2018). In this case, predicted employment growth is $\Delta\hat{\gamma_{ct}} = Oil_{ct-5} \cdot \Delta P_t$ where Oil_{ct-5} is the oil and gas employment share in CD c in year $t-5$ and ΔP_t is the national change in either oil and gas employment or the national change in the price of oil between $t-5$ and t. As in Black et al. (2002) and Charles et al. (2018), to account for sluggish local business adjustments following international price shocks, we include two lags of the instrument in ΔP when looking at oil prices.

Goldsmith-Pinkham et al. (2020) show that two identifying assumptions need to hold for consistency of the Bartik shift-share instrument: relevance and exogeneity. The relevance assumption is uncontroversial, as the first-stage coefficient of the instrument's impact on lifetime income is highly significant with F-statistic well above conventional thresholds. The exogeneity assumption requires that the industry shares must be exogenous to changes in the error term in equation (11). We test the plausibility of the exogeneity assumption by correlating CD industry shares in the initial period with initial CD characteristics. Finding a correlation between industry shares and characteristics that predict changes in the outcome variable points to omitted variables biasing estimation. As suggested by Goldsmith-Pinkham et al. (2020), we focus on the industries with the largest Rotemberg weights as well as the oil and gas sector.

Using local labor market shocks from variation in local industry composition creates three challenges for estimating the insurance value in equation (9). First, we would like to identify the pure wage effects on disability claims, but local labor market shocks may also impact job separations and job findings. Appendix A.3 shows that if labor market shocks also impact job separations and findings, we get a lower bound of the insurance value if we scale the DI inflow effect by the effect of the shock on income (the sum of earnings and unemployment insurance benefits).

Second, local labor market shocks create *temporary* variation in wages, while the 1987 reform induces a *permanent* increase in DI benefits. To compare the effects of a temporary versus a permanent change, Appendix A.4 extends the static model to a dynamic model. The model illustrates that lifetime income (instead of annual income) captures the permanent impact of a temporary wage shock. We measure average lifetime income in year t by summing up disposable income—earnings and unemployment benefits—from year t until age 60 using a discount rate of 3 percent and dividing by the number of years between year t and age 60.

Third, our model assumes that marginal entrants for changes in benefits and changes in wages are comparable. Otherwise, differential take-up responses to benefits and wages could reflect a compositional effect and might not accurately reflect differences in marginal utility between states. The empirical concern is that local economic shocks may affect different people than those affected by changes in DI benefits. To address this concern, we perform a complier analysis and compare the characteristics of the marginal types responding to local economic shocks and changes in DI benefits.

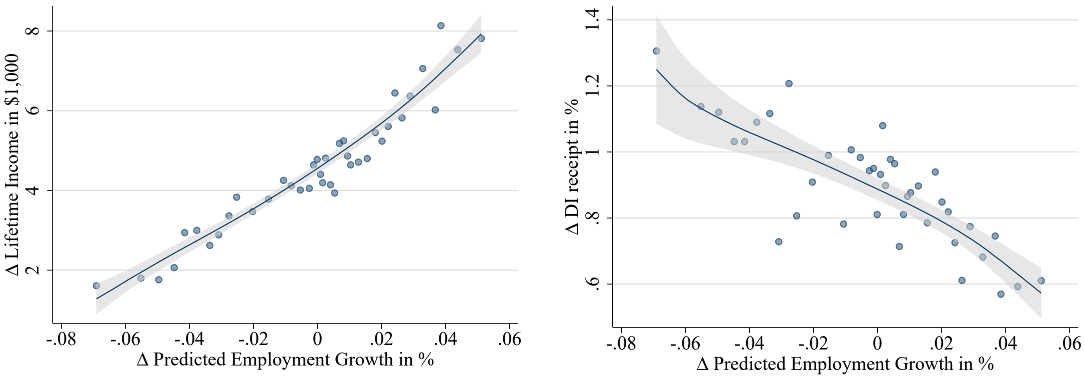
5.2 Empirical Results

Figure 4 visually illustrates our estimation strategy. The left panel presents the visual first stage, obtained from a local linear regression of the change in lifetime income in \$1,000 against the predicted employment growth from the industry shift-share instrument (the shaded area indicates a 95 percent confidence interval). The change in lifetime income is monotonically increasing in the predicted employment growth and close to linear. The right panel plots the reduced-form effect of predicted employment growth on DI receipt. DI receipt is monotonically decreasing in predicted employment growth.

Table 2 presents first-stage estimates and the OLS and 2SLS estimates of equation (11) for different earnings and income measures. Our OLS estimates show a negative impact of earnings and (lifetime) income on disability claiming. The effects are quantiatively small and align with findings by Autor and Duggan (2003) and Charles et al. (2018) for the US. The 2SLS estimates are significantly larger in magnitude. The second column shows that a \$1,000 increase in earnings reduces DI enrollment by 0.044 percentage points. We find a similar effect for current income (column 4). In contrast, the impact of a \$1,000 increase in lifetime income on DI claiming is about twice as big. This pattern is consistent with local labor market shocks having a substantial impact on current earnings and income but a smaller impact on lifetime income be-

Figure 4: Lifetime Earnings and DI Receipt by Predicted Employment Growth

(a) First Stage—Lifetime Income (b) Reduced Form—DI Receipt



Notes: The figure plots the effects of predicted employment growth rates at the census division on lifetime income (panel a) and DI receipt (panel b).

cause the shocks are temporary. Indeed, the first-stage coefficient for lifetime income is about half as large as for current earnings and income. The heteroscedasticity-robust F-statistic (Montiel Olea and Pfluger, 2013) testing the significance of the excluded instruments in the first-stage equation is 50 or larger, well above conventional thresholds.[17]

Table 3 shows that positive and negative local lifetime income shocks have similar effects on DI take-up (–0.104 vs. –0.090, respectively), and we cannot reject the hypothesis that these estimates are identical (p-value = 0.55). This suggests that DI take-up responds similarly to positive and negative income shocks.

Appendix Table C.3 reports the first-stage and 2SLS estimates for the oil instruments. The first-stage relationship between the instruments and different income measures is strong. Consistent with the shift-share estimates in Table 2, we find that higher earnings and income reduce DI take-up significantly, but the magnitude of the effect is slightly smaller. A $1,000 increase in lifetime income reduces DI take-up between 0.061 and 0.072 percentage points. However, we cannot reject the hypothesis that the industry shift-share and oil and gas 2SLS estimates are identical.[18] Appendix Table C.2

[17] Montiel Olea and Pflueger (2013) find that the critical values of the F-statistic to a test of IV relative bias not exceeding 10 percent with a significance level of 5 is between 11 and 23.1, analogous to the Stock and Yogo (2005) rule-of-thumb cutoff of 10.

[18] The p-values that the oil price and oil employment 2SLS estimates equal the Bartik 2SLS estimate are 0.284 and 0.132.

Table 2: The Impact of Earnings/Income on Disability Insurance Enrollment

	Earnings (in $1,000)		Current income (in $1,000)		Lifetime income (in $1,000)	
	OLS	2SLS	OLS	2SLS	OLS	2SLS
Δ DI take-up	-0.004⋆⋆⋆	-0.044⋆⋆⋆	-0.007⋆⋆⋆	-0.044⋆⋆⋆	-0.015⋆⋆⋆	-0.096⋆⋆⋆
	(0.001)	(0.008)	(0.002)	(0.008)	(0.003)	(0.018)
1st-stage estimate		117.0⋆⋆⋆		116.0⋆⋆⋆		53.0⋆⋆⋆
		(15.6)		(15.6)		(7.7)
Effective F-statistic		55.9		55.2		47.3
Obs.	18,829,205		18,829,205		18,829,205	

Notes: The table reports the OLS and 2SLS estimates of equation (11) for different earnings and income measures, instrumenting for ΔE_{ict} using predicted employment based on industry employment shares. Coefficient estimates and standard errors are multiplied by 100 and should be interpreted as percentage points. Standard errors are reported in parentheses and are clustered at the Census Division (CD). The sample includes 15-59 year-old individuals between 1982 and 2016. Levels of significance: ⋆10%, ⋆⋆5%, and ⋆⋆⋆1%.

Table 3: The Impact of Positive and Negative Wage Shocks on Disability Insurance Enrollment

	Earnings	Current income	Lifetime income
Negative shock, 2SLS	-0.053***	-0.054***	-0.104***
	(0.013)	(0.013)	(0.027)
Positive shock, 2SLS	-0.038***	-0.038***	-0.090***
	(0.007)	(0.007)	(0.016)
P-value, negative = positive	0.185	0.178	0.550
Negative shock, first stage	95.8***	94.4***	47.7***
	(12.1)	(12.0)	(6.2)
Positive shock, first stage	188.0***	186.7***	79.6***
	(23.3)	(23.3)	(11.3)
Negative shock, effective F-statistic	62.4	61.5	57.8
Positive shock, effective F-statistic	65.4	64.7	50.0
No. Observations		18,829,205	

Notes: The table reports 2SLS and first-stage estimates for negative and positive wage shocks using the industry shift-share instrument. Coefficient estimates and standard errors are multiplied by 100 and should be interpreted as percentage points. Standard errors are reported in parentheses and are clustered at the Census Division (CD). P-values test equality of the negative-shock and positive-shock 2SLS estimates. The sample includes 15-59 year-old individuals between 1982 and 2016. Levels of significance: ⋆10%, ⋆⋆5%, and ⋆⋆⋆1%.

documents the reduced-form estimates for the industry shift-share instrument and the impact on other outcomes. The estimate is unchanged if we set DI benefits receipt to zero for spouses who die within a year of benefit take-up (-0.095). We also find that an increase in lifetime income reduced DI payments, but has no impact on the probability to move to another CD.

The central identifying assumption for the Bartik shift-share instruments is that industry and oil employment shares are uncorrelated with shocks leading to changes in DI take-up. To assess the validity of this assumption, we analyze whether CD characteristics that could be correlated with changes in DI take-up also predict industry and oil employment shares. Appendix Table C.5 shows the relationship between initial-period CD characteristics and the CD share of the three industries with the largest Rotemberg weights (technical services, farming, and transportation), as well as the oil employment share. These characteristics explain a significant portion of the variation in industry shares across CDs: the R^2 varies between 25 and 89 percent. Reassuringly, this test reveals no systematic correlations between a CD's industry share and a CD's age composition, earnings and income distribution, mortality, and other socio-economic characteristics.[19] Appendix Figure C.2 complements this analysis by showing the full distribution of Rotemberg weights. The three largest weights sum to 0.391, or about 35 percent of total positive Rotemberg weights, while negative weights are modest: positive weights sum to 1.116 and negative weights sum to -0.116. Thus, while the largest industries account for a meaningful share of the identifying variation, the concentration is not unusually high relative to the applications in Goldsmith-Pinkham et al. (2020), and negative weights play only a limited role.

We also report alternative estimators and overidentification tests in Appendix Table C.6. The estimates are close in magnitude, suggesting that the main result is not

[19] A related concern is that application costs may vary across locations in a way that is correlated with industry composition. We cannot observe distance to the local offices where individuals could obtain information about, or submit materials for, DI applications. However, our DI take-up estimates are very similar for 1982-2001 and for the full period through 2016 (Tables 2 and C.4), even though distance to a local office should matter less once online application channels became available. This suggests that geographic variation in application costs is unlikely to drive the results.

driven by the particular weighting matrix implicit in the Bartik instrument. The main Bartik 2SLS estimate is -0.096, while the many-instrument 2SLS, MBTSLS, and LIML estimates are -0.088, -0.092, and -0.114, respectively. The overidentification tests reject the null of common effects across all industry-share instruments. As Goldsmith-Pinkham et al. (2020) emphasize, such rejections can reflect either misspecification or heterogeneous effects across instruments. To assess this directly, Appendix Figure C.3 plots each industry's just-identified 2SLS estimate against its first-stage F-statistic, with marker size scaled by the absolute Rotemberg weight. The figure shows heterogeneity across high-powered industries (i.e., those with an F-statistic above 5), but the industries with larger Rotemberg weights and stronger first stages tend to deliver estimates close to the overall Bartik estimate, while more extreme estimates receive smaller weights. We therefore interpret the overidentification-test rejection as primarily reflecting heterogeneity in industry-specific estimates rather than evidence that the main identifying variation comes from misspecified high-weight instruments.

Complier Analysis. We perform a complier analysis to compare characteristics of entrants who are marginal to changes in benefits versus changes in wages. We cannot identify individual marginal entrants, but it is possible to describe their observable characteristics. Following Frandsen et al. (2023), who extend Abadie's (2003) method for estimating complier characteristics to non-binary instruments, we run the following 2SLS specification in first differences:

$$(X \cdot \Delta DI)_{it} = \alpha + \beta \Delta DI_{it} + \pi_t + \varepsilon_{it}, \tag{13}$$

where X is an exogenous characteristic, ΔDI is a dummy for DI entry in year t and π_t are year dummies. For the benefit change, we instrument ΔDI on the right-hand side of equation (13) with $RoC_p \cdot A87_t$, the interaction of a dummy for living in the Rest of Canada and a dummy for the year 1987 or greater. This interaction captures the quasi-experimental variation in DI benefits from the 1987 reform. For the wage change, we instrument ΔDI with the industry shift-share instrument $\Delta \hat{\gamma}_{ct}$. Frandsen et al. (2023) show that β yields a weighted average of characteristic X among marginal entrants who take up DI because of a benefit or wage change.

Table 4 reports estimates of equation (13) for different characteristics. The p-value in the last column tests the null hypothesis that marginal entrant characteristics are identical for a benefit and a wage change. The first three rows show that marginal entrants of benefit and wage changes are similar in the average age, family size, and share female. The minor differences between the two groups in these characteristics are not statistically significant (p-values of 0.269, 0.132, and 0.807). Marginal entrants are older (average age around 50), live with an average of 2.6 additional family members (implying an average family size of 3.6), and are more likely to be male (female share of about 40%). The fourth row looks at the probability of dying by age 60, a proxy for disability severity θ. We find that marginal entrants' mortality rates by age 60 are comparable (13% for a benefit change and 12% for a wage change). We cannot reject the null hypothesis that the coefficients are identical (p=0.612). Moreover, we find that the share of marginal entrants across different earnings and income distribution quartiles is similar, where earnings and income is measured in $t-3$. Marginal entrants are over-represented in lower earnings quartiles: about 40% of marginal entrants come from the lowest quartile, 20-25% from the middle quartiles, and around 10% from the highest quartile.

Overall, these results indicate that our empirical variation in DI benefits and wages affects similar types. We thus identify meaningful differences in marginal utilities, and not composition effects for our main specification using the industry shift-share instrument.[20]

In the complier analysis for the oil and gas employment-share instrument, reported in Appendix Table C.8, we find that compliers differ from those identified by the benefit change along some dimensions. In particular, compliers of the oil and gas instrument have smaller families, are more likely to be female, exhibit lower mortality rates, and are drawn from somewhat higher parts of the income distribution. These differences

[20] Appendix A.2 discusses this point more formally. Equation (A.4) illustrates that we want the same (w, A, b, θ)-marginal types for both benefits and wage changes. Table 4 indicates that the empirical variation in DI benefits and wages that we exploit affect similar (w, A, θ)-types. Since DI benefits are a deterministic function of past earnings, we have comparable (w, A, b, θ)-types for benefit and wages changes.

Table 4: Comparison of Complier Characteristics

	Benefit change		Wage change		Equality test
	Coeff.	Std. err.	Coeff.	Std. err.	(P-value)
Age	50.57***	0.81	53.12***	2.15	0.269
Family size	3.60***	0.26	3.14***	0.17	0.132
Share female	0.426***	0.038	0.409***	0.054	0.807
Pr(died by age 60)	0.132***	0.017	0.119***	0.019	0.612
Share earnings 1. quartile	0.377***	0.047	0.441***	0.058	0.392
Share earnings 2. quartile	0.274***	0.026	0.256***	0.050	0.754
Share earnings 3. quartile	0.213***	0.032	0.238***	0.034	0.598
Share earnings 4. quartile	0.136***	0.034	0.065	0.063	0.320
Share income 1. quartile	0.357***	0.047	0.345***	0.040	0.849
Share income 2. quartile	0.268***	0.030	0.292***	0.053	0.703
Share income 3. quartile	0.223***	0.025	0.264***	0.030	0.285
Share income 4. quartile	0.152***	0.033	0.099	0.062	0.450

Notes: The table reports the complier characteristics (β-estimates of equation (13)) for benefit and wage changes, where the change in DI receipt on the right-hand side is instrumented with $RoC_p \cdot A87_t$ for the benefit change and industry shift-share instrument for the wage change. Earnings and income are measured in $t-3$. The last column reports the p-value for the null hypothesis that marginal entrant characteristics are identical for benefit and wage changes. Levels of significance: *10%, **5%, and ***1%.

suggest that the insurance value estimates based on the oil and gas instrument may partly reflect differences in the underlying complier populations, a point we revisit in Section 6.

6 Welfare Effects

This section estimates the insurance value and welfare impacts of changes in DI benefits. For our Canadian implementation, we combine the estimates from Sections 4 and 5. Additionally, we provide a back-of-the-envelope implementation for the U.S. DI system using estimates from the literature.

Canadian Implementation. Panel A of Table 5 reprints the DI take-up effects of a $1,000 increase in DI benefits and a $1,000 increase in lifetime income for the different instruments.[21] The first column in Panel B measures the fiscal multiplier of increasing

[21] The DI enrollment coefficient in Table 5 is scaled to represent the effect of a $1,000 increase in DI benefits. To obtain this estimate, we discount the reform-induced lump-sum increase in DI benefits to age 60 using a 3 percent discount rate and convert it into an equivalent annual benefit increase of $2,534. The coefficient in Table 5 is then calculated by dividing the reduced-form DI enrollment estimate from Table 1 (0.527)

Table 5: Welfare Analysis

A. Δ DI take-up per $1,000				
	DI benefits ($\partial DI/\partial b$)	Lifetime income ($\partial DI/\partial w$)		
		Industry share	Oil employment	Oil price
Coeff. estimate	0.208***	-0.096***	-0.061***	-0.072***
	(0.024)	(0.018)	(0.015)	(0.007)

B. Welfare impacts				
	Multiplier $\left(1+\frac{\varepsilon_{DI,b}}{1-DI}\right)$	Insurance value $\left(-\frac{\partial DI/\partial b}{\partial DI/\partial w}\right)$		
		Industry share	Oil employment	Oil price
Estimate	1.591***	2.166***	3.421***	2.885***
	(0.069)	(0.476)	(0.923)	(0.438)
P-value: estimate = 1	0.000	0.014	0.009	0.000
P-value: multiplier = ins. value		0.195	0.041	0.001

Notes: The table combines the estimates from Sections 4 and 5 to construct the multiplier and insurance value of higher DI benefits. Levels of significance: *10%, **5%, and ***1%.

DI benefits by one dollar. The fiscal multiplier is the sum of the direct costs—an additional dollar to current beneficiaries— plus indirect costs from more people entering the DI program. The indirect cost is captured by the elasticity of DI take-up with respect to benefits, normalized by the share of people not on DI. We estimate that the fiscal multiplier of an additional dollar in DI benefits is 1.6 dollars.[22]

The last three columns of Panel B report the insurance value of a one-dollar increase in DI benefits, obtained by dividing the DI take-up response to a change in DI benefits by the take-up response to a change in lifetime income. In our preferred specification using the industry shift-share instrument, the estimated insurance value is 2.2. Thus, the marginal utility of an additional dollar in the disabled state is 2.2 times the marginal

by the annual benefit increase measured in thousands of dollars, yielding $\frac{0.527}{2.534} = 0.208$.

[22] We also estimate the fiscal costs of the reform and the associated multiplier directly in reduced form in Appendix Figure B.3. Since we cannot reliably measure other benefit receipt in our data before 1992, we define fiscal costs as DI benefit expenditures minus labor income tax revenues. The implied multiplier from this exercise is 1.449, slightly below the multiplier implied by the take-up elasticity.

utility of an additional dollar in the non-disabled state.

The oil and gas instrument yields higher insurance values, ranging from 2.8 to 3.4. However, the compliers for this instrument appear wealthier and healthier than those for the benefit change (Table C.8), suggesting that they have a lower marginal utility of consumption. These estimates may therefore partly reflect differences between complier populations and may be upward biased.[23]

We can reject perfect insurance, an insurance value of one, with high confidence, as indicated by p-values around and below 1% in Panel B of Table 5. However, when we test whether the insurance value equals the multiplier, we lack the precision for the shift-share design to reject the null hypothesis that Canadian DI benefits are optimal. We find p-values of 0.195 for the industry shift-share design and 0.041 and 0.001 for the oil specifications.

As discussed in Appendix A.3, our estimator may be upward biased in the presence of benefit substitution. Table 1 implies that inflows from non-employment amount to 0.027 per $1,000, or about 13 percent of the total DI inflow effect (0.208 in Table 5). Using the bias formula from Appendix A.3, this corresponds to a bias of 0.28 in our insurance value estimate.[24] This might be interpreted as an upper bound, as it assumes that all individuals classified as non-employed are substituting other welfare benefits for DI. In practice, some may still view work as the relevant outside option. Even under

[23]A potential concern with our implementation is that the benefit effects are identified using the 1987 reform, whereas the wage effects are estimated over the longer 1982–2016 period. To address this concern, we restrict the wage analysis to 1982–2001, aligning it more closely with the timing of the benefit reform. Panel A of Appendix Table C.4 shows that limiting the sample to 1982–2001 yields estimates similar to the full-sample results. Replicating the welfare analysis for this shorter period in Appendix Table D.12 produces a slightly higher insurance value of 2.54, which is not statistically different from our baseline estimate. Overall, our main conclusions are unchanged when using the shorter sample period.

Relatedly, the oil and gas instruments exploit variation across provinces with oil and gas production. As a robustness check, we restrict both the benefit and wage analyses to the oil- and gas-producing provinces and Quebec. Panel B of Appendix Table B.2 and Appendix Table C.4 report the corresponding benefit- and wage change estimates, while Appendix Table D.13 presents the welfare analysis. The estimates remain similar, and our main conclusions are unchanged in this restricted sample.

[24]The bias is calculated as $\frac{\partial DI/\partial b|\text{from other benefits}}{-\partial DI/\partial w} = \frac{0.027}{0.096} = 0.28$.

this adjustment, the corrected insurance value is 1.88, which continues to exceed the fiscal costs.

Together, our empirical estimates indicate that DI benefits are valuable despite the relatively high costs of providing them. Because our approach likely provides a lower bound on the DI program's insurance value, as discussed in Section 2, we interpret our results as strong evidence that Canadian DI benefits are not too generous.[25]

U.S. Implementation. Milligan and Schirle (2019) estimate regressions of the form $DI_{it} = \alpha + \beta_1 \cdot \ln(B_{it}) + \beta_2 \cdot \ln(E_{it}) + \epsilon_{it}$ to identify the effect of benefit generosity, B_{it}, and earnings, E_{it}, on DI take-up, DI_{it}. They instrument for B_{it} using a simulated benefits approach and employ a Bartik shift-share design for E_{it}. We apply our sufficient statistics approach to their estimates to derive the implied insurance value and incentive costs of the U.S. DI program.

The DI take-up elasticity is $\varepsilon \equiv \frac{dDI/DI}{dB/B} = \frac{\beta_1}{DI}$. The estimates from Table 3, Column (4) in Milligan and Schirle (2019) for β_1 and the average DI rate from Table 1, Column (5) imply an elasticity of

$$\varepsilon = \frac{0.042}{0.037} = 1.14. \tag{14}$$

The fiscal multiplier is therefore $1 + \frac{\varepsilon}{1-Y} = 1 + \frac{1.14}{1-0.037} = 2.18$, i.e., a \$1 increase in U.S. DI benefits costs \$2.18.

To identify the insurance value with our approach we need $-\frac{\partial DI/\partial B}{\partial DI/\partial E}$. Table 3, Column (4) in Milligan and Schirle (2019) identifies the effect of a 100%-change in DI benefits and wages. We thus rescale the point estimates by the average DI benefit and average earnings from Table 1, Column (5): $\frac{\partial DI}{\partial B} = \frac{\beta_1}{B} = \frac{0.042}{15596}$ and $\frac{\partial DI}{\partial E} = \frac{\beta_2}{E} = \frac{-0.038}{48594}$.

[25] Our estimates are identified from the 1987 reform and therefore apply most directly to marginal beneficiaries around that period. The composition of CPP-D beneficiaries has changed since then. Women account for a larger share of beneficiaries, beneficiaries are on average older, and mental-health-related conditions represent a larger share of awards, while musculoskeletal conditions have declined somewhat in relative importance (Employment and Social Development Canada, 2025a,c,b). Whether these changes translate into different insurance values or incentive costs remains an open empirical question, as there is limited evidence on how insurance values and incentive costs vary across beneficiary characteristics.

The implied insurance value is therefore

$$\text{Insurance Value} = -\frac{\partial DI/\partial B}{\partial DI/\partial E} = 3.44, \tag{15}$$

i.e., an extra dollar in DI benefits is valued at $3.44, which exceeds the $2.18 costs of providing the extra dollar.

7 Conclusion

This paper develops and implements a novel revealed-preference approach to estimating the value of DI. Evidence on the insurance value of DI programs is limited. Existing methods to estimate the insurance value of DI programs require consumption data or data on private insurance purchases, both of which are scarce. Instead, our approach focuses on DI take-up decisions, which are observable in standard administrative data.

We show that the take-up response to a change in disability benefits relative to a change in wages identifies the implied value of disability benefits in the classic DI model from Diamond and Sheshinski (1995). In models with richer heterogeneity, our method identifies the insurance value of marginal entrants, which we argue provides a lower bound for the average insurance value of DI benefits.

In our Canadian implementation, we find significant DI take-up effects for both higher DI benefit levels and lower wages. However, the responses to benefit changes are quantitatively larger. Combining the wage and DI benefit estimates, we find that the insurance value of a $1 increase in DI benefits is $2.2 when exploiting differential exposure to industry shocks across census divisions. We estimate even higher insurance values of $2.8 to $3.4 for oil price shocks. For the incentive costs of higher DI benefits, we find that an extra dollar in disability benefits costs $1.6 in total. Our estimates indicate that the insurance value is large relative to the fiscal costs and suggest that Canadian DI benefits are not too generous. Similarly, U.S. DI benefits are not too generous based on applying our approach to existing U.S. estimates from Milligan and Schirle (2019).

While we focus on the insurance value of the Canadian and US DI programs, our theoretical framework is more general. The method can be applied in countries where data on take-up decisions are available and it would be valuable to know whether the same conclusions arise in other DI systems. The same logic also applies to other social insurance programs in which individuals choose between working and program take-up, and in which work and benefit receipt are mutually exclusive. Examples include old-age pensions, sick leave, and disability insurance programs. Applying our framework in these settings could deepen our understanding of the incentive-insurance trade-off across social insurance programs.

References

Abadie, A. (2003). Semiparametric Instrumental Variable Estimation of Treatment Response Models. *Journal of Econometrics 113*(2), 231–263.

Autor, D. H. and M. G. Duggan (2003). The Rise in the Disability Rolls and the Decline in Unemployment. *Quarterly Journal of Economics 118*(1), 157–206.

Baker, M. and D. Benjamin (1999). Early Retirement Provisions and the Labor Force Behavior of Older Men: Evidence from Canada. *Journal of Labor Economics 17*(4), 724–756.

Baker, M. and K. Milligan (2012). Disability insurance programs in Canada. In D. A. Wise (Ed.), *Social security programs and retirement around the world: Historical trends in mortality and health, employment, and disability insurance participation and reforms*, pp. 327–58. Chicago:University of Chicago Press.

Black, D., K. Daniel, and S. Sanders (2002). The Impact of Economic Conditions on Participation in Disability Programs: Evidence from the Coal Boom and Bust. *American Economic Review 92*(1), 27–50.

Bordalo, P., N. Gennaioli, and A. Shleifer (2022). Overreaction and diagnostic expectations in macroeconomics. *Journal of Economic Perspectives 36*(3), 223–244.

Bound, J. and R. V. Burkhauser (1999). Economic Analysis of Transfer Programs Targeted on People with Disabilities. In *Handbook of Labor Economics*, Volume 3, Chapter 51, pp. 3417–3528. Elsevier.

Brown, J. R., G. S. Goda, and K. McGarry (2016, April). Heterogeneity in state-dependent utility: Evidence from strategic surveys. *Economic Inquiry 54*(2), 847–861.

Cabral, M. and M. R. Cullen (2019, August). Estimating the value of public insurance using complementary private insurance. *American Economic Journal: Economic Policy 11*(3), 88–129.

Campolieti, M. and C. Riddell (2012). Disability Policy and the Labor Market: Evidence from a Natural Experiment in Canada, 1998–2006. *Journal of Public Economics 96*(3-4), 306–16.

Charles, K. K., Y. Li, and S. J. Melvin (2018). Disability Benefit Take-Up and Local Labor Market Conditions. *The Review of Economics and Statistics*.

Chetty, R. (2006). A General Formula for the Optimal Level of Social Insurance. *Journal of Public Economics 90*(10-11), 1879–1901.

Chetty, R. (2008). Moral Hazard versus Liquidity and Optimal Unemployment Insurance. *Journal of Political Economy 116*(2), 173–234.

Clausen, A. and C. Strub (2020). Reverse Calculus and Nested Optimization. *Journal of Economic Theory 187*, 105019.

D'Acunto, F., M. Weber, and X. Yin (2024). Subjective income expectations and household debt cycles. NBER Working Paper 32715, National Bureau of Economic Research.

Deshpande, M. and L. M. Lockwood (2022). Beyond health: Nonhealth risk and the value of disability insurance. *Econometrica 90*(4), 1781–1810.

Diamond, P. A. and E. Sheshinski (1995). Economic Aspects of Optimal Disability Benefits. *Journal of Public Economics 57*(1), 1–23.

Diarra, M., G. Pelletier, and M. Pépin (2015). Évolution de la clientèle de la rente d'invalidité de 1970 à 2010. Technical report, Régie des Rentes du Québec.

Eisenhauer, P., J. J. Heckman, and E. Vytlacil (2015). The generalized roy model and the cost-benefit analysis of social programs. *Journal of Political Economy 123*(2), 413–443.

Employment and Social Development Canada (2018). 2017-18 Departmental Results Report: Supporting Information on Lower-Level Programs. Technical report, Government of Canada.

Employment and Social Development Canada (2025a). Canada Pension Plan (CPP) - Disability Benefits by Class of Diagnosis and Age Group. Technical report, Government of Canada Open Data.

Employment and Social Development Canada (2025b). Canada Pension Plan (CPP) - Number of New Disability Benefits by Age, Gender and by Calendar Year. Technical report, Government of Canada Open Data.

Employment and Social Development Canada (2025c). Monthly Average Number of

Canada Pension Plan (CPP) Disability Benefits by Province, Gender, and Calendar Year. Technical report, Government of Canada Open Data.

Finkelstein, A., E. F. Luttmer, and M. J. Notowidigdo (2013, January). What Good Is Wealth Without Health? The Effect of Health on the Marginal Utility of Consumption. *Journal of the European Economic Association 11*(1), 221–258.

Frandsen, B., L. Lefgren, and E. Leslie (2023). Judging Judge Fixed Effects. *American Economics Review 113*(1), 253–77.

Goldsmith-Pinkham, P., I. Sorkin, and H. Swift (2020). Barik Instruments: What, When, Why and How. *American Economic Review 110*(8), 2586–2624.

Gruber, J. (1997). The Consumption Smoothing Benefits of Unemployment Insurance. *American Economic Review 87*(1), 192–205.

Gruber, J. (2000). Disability Insurance Benefits and Labor Supply. *Journal of Political Economy 108*(6), 1162–1183.

Hendren, N. (2017, July). Knowledge of future job loss and implications for unemployment insurance. *American Economic Review 107*(7), 1778–1823.

Landais, C. (2015). Assessing the Welfare Effects of Unemployment Benefits Using the Regression Kink Design. *American Economic Journal: Economic Policy 7*(4), 243–78.

Landais, C. and J. Spinnewijn (2021). The Value of Unemployment Insurance. *The Review of Economic Studies 88*(6), 3041–3085.

Lillard, L. A. and Y. Weiss (1997, April). Uncertain health and survival: Effects on end-of-life consumption. *Journal of Business & Economic Statistics 15*(2), 254–268.

Low, H. and L. Pistaferri (2015). Disability Insurance and the Dynamics of the Incentive Insurance Trade-Off. *American Economic Review 105*(10), 2986–3029.

Low, H. and L. Pistaferri (2020). Disability Insurance: Theoretical Trade-Offs and Empirical Evidence. *Fiscal Studies 41*(1), 129–164.

Maestas, N., K. J. Mullen, and A. Strand (2013). Does Disability Insurance Receipt Discourage Work? Using Examiner Assignment to Estimate Causal Effects of SSDI Receipt. *American Economic Review 103*(5), 1797–1829.

Meyer, B. D. and W. K. Mok (2019). Disability, Earnings, Income and Consumption. *Journal of Public Economics 171*, 51–69.

Milligan, K. and T. Schirle (2019). Push and Pull: Disability Insurance, Regional Labor Markets, and Benefit Generosity in Canada and the United States. *Journal of Labor Economics 37*(S2), S289–S323.

Montiel Olea, J. and C. Pflueger (2013). A Robust Test for Weak Instruments. *Journal of Business and Economic Statistics 31*(3), 358–369.

Québec, R. (2024). Rapport annuel de gestion 2023. Technical report.

Rozsypal, F. and K. Schlafmann (2023). Overpersistence bias in individual income expectations and its aggregate implications. *American Economic Journal: Macroeconomics 15*(4), 331–371.

Seibold, A., S. Seitz, and S. Siegloch (2025). Privatizing disability insurance. *Econometrica 93*(5), 1697–1737.

Shimer, R. and I. Werning (2007). Reservation Wages and Unemployment Insurance. *Quarterly Journal of Economics 122*(3), 1145–1185.

Social Security Administration (2010). Disability Impairments on Cases Most Frequently Denied by Disability Determination Services and Subsequently Allowed by Administrative Law Judges. Technical report, Office of the Inspector General, Audit Report A-07-09-19083.

Social Security Tribunal of Canada (2025). Annual Report 2024-2025: Putting People at the Centre of Our Service. Technical report.

Staubli, S. and Q. Zhao (2023). The Long-Run and Distributional Impacts of Lowering the Pension Eligibility Age. *Working Paper*.

Stock, J. H. and M. Yogo (2005). Testing for Weak Instruements In Linear IV Regression. In D. W. Andrews and J. H. Stock (Eds.), *Identification and Inference for Econometric Models: Essays in Honor of Thomas Rothenberg*, pp. 80–108. Cambridge: Cambridge University Press.

Torjman, S. (2002). The Canada Pension Plan Disability Benefit. *Caledon Institute of Social Policy*.

APPENDIX

A Theoretical Framework: Extensions and Robustness

This Appendix derives the optimal benefit formula from Section 2 in Appendix A.1 and discusses three extensions of the simple model from Section 2. The first extension in Appendix A.2 introduces non-separability in the utility function and heterogeneity in wages, DI benefits, and unearned income. We show that in this case our approach identifies the insurance value of marginal entrants. Appendix A.2 then also discusses how the insurance value of the marginal entrant relates to the insurance value of the average DI recipient. The second extension in Appendix A.3 provides a model of endogenous job search with three labor market states: work, DI benefits, and other benefits. We show that we identify a lower bound on the insurance value if we identify the denominator in equation (9) from exogenous labor market shocks that shift both wages and employment probabilities. The third extension in Appendix A.4 provides a dynamic model to address the question how we can make permanent DI benefit changes comparable to temporary wage changes.

A.1 Derivation Optimal DI Benefits

Combining equations (3)-(5) we can write the welfare effect of a balanced-budget increase of DI benefits as

$$\begin{aligned}\frac{dW(b)}{db} &= -\int_0^{\theta^A}\left[u'(w-\tau)\frac{d\tau}{db}\right]dF(\theta)+\int_{\theta^A}^{\infty}\left[p(\theta)v'(b)-(1-p(\theta))(u'(w-\tau)\frac{d\tau}{db})\right]dF(\theta)\\ &= -\left[1-DI\right]u'(w-\tau)\frac{d\tau}{db}+DI\cdot v'(b) \qquad \text{(A.1)}\end{aligned}$$

where $\frac{d\tau}{db} = \frac{DI}{1-DI}\left(1+\frac{dDI}{db}\frac{b}{DI}\frac{1}{1-DI}\right) = \frac{DI}{1-DI}\left(1+\frac{\varepsilon_{DI,b}}{1-DI}\right)$ following from the balanced budget constraint and $\frac{dDI}{db} = \frac{d\theta^A}{db}p(\theta^A)f(\theta^A)$. Because of the envelope theorem changes in DI take-up ($\frac{d\theta^A}{db}$) have no first-order welfare effects and only enter through the government budget constraint ($\frac{d\tau}{db}$). Setting $\frac{dW(b)}{db} = 0$ for the optimal benefit level and rearranging yields the classic Baily-Chetty formula in equation (6).

A.2 Non-Separability and Heterogeneity

The two critical assumptions that the relative take-up responses identify the insurance value of DI benefits are that (i) the disability level enters the utility function additively and (ii) individuals only differ in their disability level. The two assumptions ensure that the insurance values of the marginal entrant and the average DI recipient are the same. We relax both assumptions here, showing that with non-separability or richer heterogeneity, our approach identifies the insurance value for marginal entrants.

Non-Separability of Consumption and Disutility of Work. The classic DI model in Diamond and Sheshinski (1995) assumes that disability (or disutility of work) enters the utility function additively: utility while working is $u(c^w)-\theta$ in Diamond and Sheshinski (1995). More general utility functions would be of the form $u(c^w, \theta)$ and $v(c^b, \theta)$. The application decision is then: apply if $v(c^b, \theta) - u(c^w, \theta) \geq 0$. As long as there is monotonicity, i.e., more disabled individuals are more likely to apply, there is again a marginal entrant determined by the solution to $H = v(c^b, \theta^A) - u(c^w, \theta^A) = 0$. We can then apply the implicit function theorem, and the relative DI take-up effects are given by

$$-\frac{\partial DI}{\partial b} / \frac{\partial DI}{\partial w} = \frac{v'(c^b, \theta^A)}{u'(c^w, \theta^A)}. \tag{A.2}$$

Thus, our approach still measures the insurance value of marginal entrants, but it may no longer be representative for the average DI recipient's insurance value.

Heterogeneity. To introduce heterogeneity beyond disability levels, we let after-tax wages $w_i - \tau_i$, DI benefits b_i, unearned income A_i, and disability level θ_i vary across individuals i. Consumption of individual i when working is then given by $c_i^w = w_i - \tau_i + A_i$ and consumption when on DI benefits is $c_i^b = b_i + A_i$. Assume there is a distribution of wages, DI benefits, and unearned income $G(w, b, A)$ and a conditional distribution of disability $F(\theta | w, b, A)$. The marginal entrant for a given wage, DI benefit, and unearned income is given by

$$\theta^A(w,b,A) = u(c^w) - v(c^b). \tag{A.3}$$

DI take-up is determined by $DI = \int \int_{\theta^A(w,b,A)}^{\infty} p(\theta) dF(\theta|w,b,A) dG(w,b,A)$. Therefore,

$$-\frac{\partial DI}{\partial b} / \frac{\partial DI}{\partial w} = \frac{\int \left[v'(c^b) f(\theta^A) p(\theta^A) db \right] dG(w,b,A)}{\int \left[u'(c^w) f(\theta^A) p(\theta^A) dw \right] dG(w,b,A)}. \tag{A.4}$$

The relative DI take-up responses with respect to DI benefits and wages still identify the insurance value of marginal entrants—the ratio of the marginal utilities of consumption on DI benefits and while working—but it could again differ from the insurance value of the average DI recipient.

Composition or Insurance Value? When implementing equation (A.4), an empirical challenge arises. The empirical variation to identify DI take-up effects of benefits and wages might not affect the same individuals. In this case, the differential take-up responses we identify empirically could be driven by composition (different people are affected) rather than by differences in marginal utilities between working and receiving DI benefits. We address this concern in Section 5 by implementing a complier analysis that compares the characteristics of marginal entrants for changes in benefits with those of marginal entrants for changes in wages. Reassuringly, Table 4 shows that the incidence of these two groups of marginal entrants across the earnings and taxable income distribution is strikingly similar. Moreover, mortality, a measure of disability severity, is not statistically different between the two complier groups. These findings indicate that the empirical variation in DI benefits and wages affect similar (w, A, θ)-types, suggesting we identify meaningful differences in marginal utilities and not composition effects.[26]

When is the Insurance Value of the Marginal Entrant a Lower Bound? A natural question is how representative is the marginal entrant's insurance value for the DI program's insurance value? What we want to measure is the value of a transfer between DI

[26]Since DI benefits are a deterministic function of past earnings, we have comparable (w, A, b, θ)-types.

and non-DI recipients, which we can write as

$$IV_{pop} = \frac{E\left[v'(c^b(\theta);\theta)|\text{on DI}\right]}{E\left[u'(c^w(\theta);\theta)|\text{not on DI}\right]}, \tag{A.5}$$

where we allow consumption and utility itself to vary by the disability level. Our approach identifies the value of a transfer between the DI and the non-DI state for marginal entrants

$$IV_A = \frac{E\left[v'(c^b(\theta);\theta)|\theta^A\right]}{E\left[u'(c^w(\theta);\theta)|\theta^A\right]}. \tag{A.6}$$

The question is whether $IV_A \leq IV_{pop}$. We discuss this question in two steps. First, we analyze the case when the marginal utility of consumption only depends on consumption and is independent of the disability level θ. Second, we discuss the case when the marginal utility of consumption varies with the disability level.

If the marginal utility of consumption is independent of the disability level—the assumption in Diamond and Sheshinski (1995)—then $IV_A \leq IV_{pop}$ holds if (i) utility functions are concave, i.e., the marginal utility of consumption falls with consumption, (ii) the marginal entrant is weakly better off than the average DI recipient when on DI benefits, $E\left[c^b(\theta)|\theta^A\right] \geq E\left[c^b(\theta)|\text{on DI}\right]$, and (iii) the marginal entrant is weakly worse off than the average taxpayer when working, $E\left[c^w(\theta)|\theta^A\right] \leq E\left[c^w(\theta)|\text{not on DI}\right]$. Under assumptions (i)-(iii) we have

$$\frac{E\left[v'(c^b(\theta))|\theta^A\right]}{E\left[u'(c^w(\theta))|\theta^A\right]} \leq \frac{E\left[v'(c^b(\theta))|\text{on DI}\right]}{E\left[u'(c^w(\theta))|\text{not on DI}\right]} \tag{A.7}$$

and the insurance value of marginal entrants indeed provides a lower bound for the true insurance value of DI benefits. Whether assumptions (ii) and (iii) hold is an empirical question. Consumption when on DI benefits is $c^b(\theta) = b(\theta) + A(\theta)$ – DI benefits plus other non-labor income. Consumption when working is $c^w(\theta) = w(\theta) + A(\theta)$ – labor income plus other non-labor income. We do not directly observe consumption in the data. However, we can characterize the taxable incomes for the marginal entrant with our complier analysis from Section 5 and compare it to the average taxpayer and the average DI recipient.

Table 4 shows that marginal entrants are more likely to come from the lower end of the income distribution. Approximately 35% of all marginal entrants are in the lowest income quartile and only 10-15 percent are in the highest income quartile. Hence, marginal entrants are over-represented at lower incomes and under-represented at higher incomes (for equal representation, 25% of marginal entrants should be in each quartile). Income is measured three years before DI take-up, i.e., it is a measured when marginal entrants are still in the labor market and includes all types of taxable incomes, including labor and capital income. These patterns indicate that marginal entrants fare worse than the average taxpayer and support assumption (iii) $E\left[c^w(\theta)|\theta^A\right] \leq E\left[c^w(\theta)|\text{not on DI}\right]$.

Table B.3 compares marginal entrants and the average DI recipient.[27] Marginal entrants are on average older and less likely to be female. Interestingly, marginal entrants have a similar mortality as the average DI recipient. Marginal and average entrant have comparable incomes three years before DI take-up. Since average and marginal entrants have a very similar incidence across the income distribution, we might expect that they fare equally well when on DI benefits, i.e., $E\left[c^b(\theta)|\theta^A\right] \approx E\left[c^b(\theta)|\text{on DI}\right]$. Taken together, our complier analysis suggests that marginal entrants are similar to the average DI recipient in terms of income and mortality and less well off than the average taxpayer.

Even if assumptions (i)-(iii) hold, the marginal entrant's insurance value might not be a lower bound for the average insurance value if the marginal utility of consumption varies with the disability level. The marginal utility of consumption can vary with the disability level in two ways. The first scenario is that the marginal utility of consumption increases with the disability level, which is the modeling assumption in the structural DI model of Low and Pistaferri (2015). In this case, $v'(c^b;\theta^A) \leq E\left[v'(c^b;\theta)|\text{on DI}\right]$ and $u'(c^w;\theta^A) \geq E\left[u'(c^w;\theta)|\text{not on DI}\right]$ as long as more disabled individuals are more likely to end up on DI benefits, i.e., if ($E\left[\theta|\text{not on DI}\right] \leq \theta^A \leq E\left[\theta|\text{on DI}\right]$ holds. Hence, the marginal entrants' insurance value provides a lower bound in this scenario.

The second scenario is that the marginal utility of consumption is lower for higher disability levels. Then, the insurance value of the marginal entrants might not be a lower

[27]Tables C.7 and C.9 provide the same comparisons for the marginal entrant of the industry shift-share and the oil and gas instruments.

bound for the actual insurance value.

Therefore, an important determinant of the lower-bound result is how marginal utility varies with health. The empirical evidence on such state dependence is inconclusive. Some studies find that marginal utility declines in worse health states (Finkelstein et al., 2013), while others find the opposite (Lillard and Weiss, 1997). Additional evidence points to limited average differences but substantial heterogeneity across individuals (Brown et al., 2016). As a result, whether the marginal entrant's insurance value is below that of the average DI recipient is ultimately an empirical question. The DI literature provides little direct evidence on this comparison. To our knowledge, no study compares the insurance value of marginal entrants with the average insurance value of the DI program, leaving this as an important area for future research.

A.3 Model with Three Labor Market States

Job Search Model. We now extend the model with a third labor market state—unemployed—and endogenous job search to address an empirical challenge. From a theoretical perspective, this extension does not change our results as long as the relevant outside option for the marginal entrant is work—an assumption we relax below. The relative DI responses to changes in benefits and wages continue to identify the insurance value. However, empirically, it is challenging to isolate pure wage effects. The shift-share instruments we use affect wages but can also influence separation and job-finding rates.

In the extended model, individuals are employed with probability $s(e;\Omega)$, which depends on individual effort e and exogenous economic conditions Ω. Individuals are unemployed with probability $1 - s(e;\Omega)$. An individual's wage, $w(e;\Omega)$, is also a function of individual effort and exogenous economic conditions. Effort is costly and creates utility loss $\psi(e;\theta)$, which varies by the disability level θ. The optimal effort level, e, is determined by

$$\max_{e} s(e;\Omega) \cdot u(w(e;\Omega)) + (1 - s(e;\Omega)) \cdot v(z) - \psi(e;\theta) \tag{A.8}$$

An interior solution fulfills the following first-order condition

$$s'(e;\Omega)\cdot[u(w(e;\Omega))-v(z)]+s(e;\Omega)\cdot u'(w;\Omega)\cdot\frac{\partial w(e;\Omega)}{\partial e}-\psi'(e;\theta)=0. \quad \text{(A.9)}$$

Individuals with too high θ might choose a corner solution with $e = 0$. In this case, we assume that $s(0) = 0$ and individuals receive other welfare benefits z with certainty.

If individuals are awarded DI, they leave the labor force permanently and do not incur any effort cost/disutility; they have flow utility $v(b)$. Absent application costs, an individual applies for DI benefits if the expected utility of receiving DI exceeds the expected utility of being in the labor force:

$$v(b)\geq s(e;\Omega)\cdot u(w)+(1-s(e;\Omega))\cdot v(z)-\psi(e;\theta) \quad \text{(A.10)}$$

The marginal entrant θ^A is determined by

$$\Theta\equiv s(e;\Omega)\cdot u(w)+(1-s(e;\Omega))\cdot v(z)-\psi(e;\theta^A)-v(b)=0. \quad \text{(A.11)}$$

From (A.11), a change in DI benefits has the following impact on application behavior:

$$\frac{\partial\theta^A}{\partial b}=\frac{v'(b)}{\partial\Theta/\partial\theta^A}. \quad \text{(A.12)}$$

An economic shock Ω has the following impact on application behavior:

$$\frac{\partial\theta^A}{\partial\Omega}=-\frac{1}{\partial\Theta/\partial\theta^A}\left[\frac{\partial s(e;\Omega)}{\partial\Omega}[u(w)-v(z)]+s\cdot u'(w)\cdot\frac{\partial w(e;\Omega)}{\partial\Omega}\right]. \quad \text{(A.13)}$$

A negative economic shock thus increases applications through two channels: (i) increased job loss $\frac{\partial s(e;\Omega)}{\partial\Omega}$ and (ii) reduced wages $\frac{\partial w(e;\Omega)}{\partial\Omega}$. The ratio of the take-up response to changes in benefits and economic conditions is given by

$$\frac{\frac{\partial\theta^A}{\partial b}}{-\frac{\partial\theta^A}{\partial\Omega}}=\frac{v'(b)}{\left[\frac{\partial s(e;\Omega)}{\partial\Omega}[u(w)-v(z)]+s\cdot u'(w)\cdot\frac{\partial w(e;\Omega)}{\partial\Omega}\right]} \quad \text{(A.14)}$$

$$\leq\frac{v'(b)}{u'(w)\left[\frac{\partial s(e;\Omega)}{\partial\Omega}[w-z]+s\cdot\frac{\partial w(e;\Omega)}{\partial\Omega}\right]}. \quad \text{(A.15)}$$

The inequality in (A.15) holds if $u'(w)(w-z) \leq u(w) - v(z)$. That is, the monetized utility loss associated with job loss, $(u(w) - v(z))/u'(w)$, is at least as large as the income loss associated with job loss, $(w - z)$. If the utility function is not state-dependent, i.e., $u(\cdot) = v(\cdot)$, and the replacement rate of other benefits is less than 100 percent, $w \geq z$, the condition holds for concave utility functions (falling marginal utility of consumption). If we assume $u'(w)(w-z) \leq u(w) - v(z)$, we can rescale the economic shock, $\partial\Omega$, by its impact on income (=earnings plus benefits) $\frac{\partial}{\partial\Omega}\left[s \cdot w + (1-s) \cdot z\right] = \left[\frac{\partial s(e;\Omega)}{\partial\Omega}\left[w - z\right] + s \cdot \frac{\partial w(e;\Omega)}{\partial\Omega}\right]$ and get

$$\frac{\frac{\partial\theta^A}{\partial b}}{-\frac{\partial\theta^A}{\partial\Omega} / \left[\frac{\partial s(e;\Omega)}{\partial\Omega}\left[w - z\right] + s \cdot \frac{\partial w(e;\Omega)}{\partial\Omega}\right]} \leq \frac{v'(b;\theta^A)}{u'(w;\theta^A)} \leq \frac{E\left[v'(b;\theta)|\text{on DI}\right]}{E\left[u'(w;\theta)|\text{working}\right]} \tag{A.16}$$

Hence, our empirical estimate that rescales the inflow effect by income is a lower bound for the insurance value of the marginal entrant.

Benefit Substitution Model. For some marginal entrants, the outside option is not necessarily work but may instead be other welfare benefits. In such cases, an increase in DI benefit generosity induces benefit substitution for a subset of marginal entrants. This extension shows that benefit substitution creates an upward bias in our estimator, as changes in DI benefits generate inflows from other welfare programs while changes in wages do not. The magnitude of this bias can be quantified empirically.

Such benefit substitution can occur in a model where individuals differ in their level of disability and application costs. We assume that there are other welfare benefits z and a DI application has cost ψ. An individual with disability severity θ and application costs ψ applies for DI benefits if their expected gain from receiving DI benefits exceeds their application costs:

$$p(\theta)\left[v(b) - \max\left\{u(w-\tau) - \theta, v(z)\right\}\right] - \psi \geq 0. \tag{A.17}$$

Individuals with high application costs might not apply for DI benefits even if they are severely disabled such that they do not work and claim other welfare benefits. In this setting there is a marginal entrant for each level of application costs, $\theta^A(\psi)$. The marginal entrants are implicitly defined by

$$p(\theta^A(\psi))\left[v(b)-\max\left\{u(w-\tau)-\theta^A(\psi),v(z)\right\}\right]-\psi=0. \tag{A.18}$$

The relevant outside option to DI benefits differs by disability severity, $\max\left\{u(w-\tau)-\theta^A(\psi),v(z)\right\}$. Let $\theta^Z=u(w-\tau)-v(z)$ denote the cutoff for the outside option where individuals with $\theta>\theta^Z$ chose other benefits over working. In this model the share of individuals on DI is given by

$$DI=\int_0^\infty\int_{\theta^A(\psi)}^\infty p(\theta)dF(\theta|\psi)dH(\psi), \tag{A.19}$$

where $F(\theta|\psi)$ is the distribution of disability severity conditional on a given level of application costs and $H(\cdot)$ denotes the distribution of application costs.

A change in DI benefit generosity then affects DI take-up by

$$\frac{\partial DI}{\partial b}=\int_0^\infty p(\theta^A)f(\theta^A|\psi)\left[1\left\{\theta^A>\theta^Z\right\}\frac{v'(b)}{\partial\tilde{F}/\partial\theta^A}+1\left\{\theta^A\le\theta^Z\right\}\frac{v'(b)}{\partial F/\partial\theta^A}\right]dH(\psi), \tag{A.20}$$

which illustrates that take-up can come from either individuals on other benefits $\theta^A>\theta^Z$ or from individuals working $\theta^A\le\theta^Z$.[28]

A change in wages affects DI take-up by

$$\frac{\partial DI}{\partial w}=-\int_0^\infty p(\theta^A)f(\theta^A|\psi)\left[1\left\{\theta^A\le\theta^Z\right\}\frac{u'(w-\tau)}{\partial F/\partial\theta^A}\right]dH(\psi) \tag{A.21}$$

and hence is only driven by individuals for whom the outside option is working.

In this setup, the ratio of take-up responses with respect to benefits and wages is

$$\begin{aligned}\frac{\partial DI/\partial b}{-\partial DI/\partial w} &= \frac{v'(b)}{u'(w-\tau)}+\frac{\int_0^\infty p(\theta^A)f(\theta^A|\psi)\left[1\left\{\theta^A\le\theta^Z\right\}\frac{v'(b)}{\partial F/\partial\theta^A}\right]dH(\psi)}{\int_0^\infty p(\theta^A)f(\theta^A|\psi)\left[1\left\{\theta^A\le\theta^Z\right\}\frac{u'(w-\tau)}{\partial F/\partial\theta^A}\right]dH(\psi)}\\ &= \frac{v'(b)}{u'(w-\tau)}+\frac{\partial DI/\partial b|\text{from other benefits}}{-\partial DI/\partial w}.\end{aligned} \tag{A.22}$$

[28] $\tilde{F}\equiv p(\theta^A(\psi))\left[v(b)-v(z)\right]-\psi$ and $F\equiv p(\theta^A(\psi))\left[v(b)-\left\{u(w-\tau)-\theta^A(\psi)\right\}\right]-\psi$

A key implication is that wage changes only affect individuals whose outside option is work, whereas DI benefit changes also affect individuals whose outside option is other welfare programs. In this setting, our revealed preference method identifies the insurance value, $\frac{v'(b)}{u'(w-\tau)}$, plus a bias term that reflects the extent to which increases in DI generosity induce inflows from other welfare programs relative to those induced by wage changes. As a result, our revealed preference method overstates the insurance value, but the magnitude of this bias can be quantified empirically by decomposing inflows into transitions from work and from other benefits.

A.4 Dynamic Model

We now extend the model to a dynamic setting. We first discuss the dynamic version of the model from the main text with two labor market states in Section A.4.1. We then discuss the dynamic version of the model from Appendix A.3 with three labor market states in Section A.4.2. The findings from the static model generalize to the dynamic setting. The challenge in the dynamic setup is empirical because we need to make temporary changes in wages comparable to permanent changes in DI benefits.

A.4.1 Dynamic Model with Two Labor Market States.

Consider the dynamic version of the simple static model from Section 2. We extend the model to T periods and allow for richer heterogeneity beyond θ. θ and other state variables evolve stochastically over the agent's relevant time horizon. Let $X_{i,t} = \{\theta_{i,t}, A_{i,t}, \chi_{i,t}\}$ denote the vector of state variables where $\theta_{i,t}$ denotes agent i's disability level in period t, $A_{i,t}$ denotes the asset level, and $\chi_{i,t}$ is a vector of other state variables (allowing for heterogeneity across agents in other dimensions like differences in education or experience that translate to differences in wages). The state vector $X_{i,t}$ summarizes all the information relevant for agent i's choices in period t. The laws of motion of assets in the disability and employment state are

$$A_{i,t+1} = (1 + r_t)A_{i,t} + b(X_{i,t}) - c^b_{i,t}(X_{i,t}) \quad \text{(A.23)}$$

$$A_{i,t+1} = (1 + r_t)A_{i,t} + w(X_{i,t}) - \tau_{i,t}(X_{i,t}) - c^w_{i,t}(X_{i,t}), \quad \text{(A.24)}$$

where DI benefits of individual i in period t, $b(X_{i,t})$, can depend on the agent's state vector $X_{i,t}$. Analogously, $w(X_{i,t})$ denotes labor income and $\tau(X_{i,t})$ are taxes. Agents make state contingent plans on how much to consume in each labor market state $\left\{c_{i,t}^{b}(X_{i,t}), c_{i,t}^{w}(X_{i,t})\right\}$, and whether they apply to DI benefits $\alpha_{i,t}(X_{i,t}) \in \{0, 1\}$.

The sequence of events and choices is as follows. At the beginning of the period, the shocks $\theta_{i,t}$ and $\chi_{i,t}$ are revealed. Having learned $X_{i,t}$, the individual decides whether to apply to DI, $\alpha_{i,t}(X_{i,t}) = 1$, or not, $\alpha_{i,t}(X_{i,t}) = 0$. If the application is accepted, individuals become DI beneficiaries in the next period, $t + 1$, for the rest of their life. If the application is rejected, individuals stay in the labor market and can apply again the next period.

Denote by $DI_{i,t+1}$ the probability that an agent i is a DI benefit recipient in period $t + 1$:

$$DI_{i,t+1} \quad = \quad 1 - \left[\prod_{k=0}^{t} \left(1 - \alpha_{i,k}(X_{i,k}) \cdot p\left(\theta_{i,k}\right)\right) \right]. \tag{A.25}$$

The probability that agent i transitions to DI in period $k + 1$ is $\alpha_{i,k}(X_{i,k}) \cdot p\left(\theta_{i,k}\right)$. The other state variables, disability level $\theta_{i,t}$ and $\chi_{i,t}$, follow stochastic processes that can, in principle, depend on agents' choices. The expectation operator $E\left[\cdot\right]$ below captures the evolution of the state variables.[29] The agent's problem is given by

$$\begin{aligned} V_i(b) \quad = \quad \max E &\left[\sum_{t=0}^{T-1} \beta^t \left(v(c_{i,t}^{b}) \cdot D_{i,t} + \left(u(c_{i,t}^{w}) - \psi(\theta_{i,t})\right) \cdot (1 - D_{i,t})\right)\right] \\ +E &\left[\sum_{t=0}^{T-1} \beta^t \mu_{i,t}^{D} \left((1 + r_t)A_{i,t} + b_{i,t} - c_{i,t}^{b} - A_{i,t+1}\right) D_{i,t}\right] \\ +E &\left[\sum_{t=0}^{T-1} \beta^t \mu_{i,t}^{W} \left((1 + r_t)A_{i,t} + w_{i,t} - \tau_{i,t} - c_{i,t}^{w} - A_{i,t+1}\right) (1 - D_{i,t})\right] \end{aligned} \tag{A.26}$$

[29]The operator $E\left[Y\right]$ aggregates the variable Y over states of nature, i.e., $E\left[Y\right] = \int Y(X_{i,t})dF(X_{i,t})$ where $F(\cdot)$ is the distribution of the state variables $X(i, t)$. This formulation is flexible: the only restriction we impose on the distribution of state variables is that it does not directly depend on DI benefits $\{b_t\}_{t=0}^{T-1}$ itself. The evolution of $X_{i,t}$, however, can depend on the agent's choices, which themselves depend on DI benefits.

Planner's Problem. The social planner maximizes social welfare by choosing DI benefits and solves

$$\max_b W(b) = \int_i V_i(b)di + \lambda \cdot G(b) \tag{A.27}$$

where

$$G(b) = \int_i E\left[\sum_{t=0}^{T-1} (1+r_t)^{-t} \left((1-D_{i,t}) \cdot \tau_{i,t} - D_{i,t} \cdot b_{i,t}\right)\right] di \tag{A.28}$$

is the planners net revenue and λ denotes the Lagrange multiplier on the planner's budget constraint.

Optimal DI Benefit. We focus on permanent and uniform DI benefits: $b(X_{i,t}) = b$ for all $X_{i,t}$. This resembles our empirical setup where the flat DI benefit uniformly changed for all DI recipients.[30] The optimal DI benefit solves

$$\frac{\partial W(b)}{\partial b} = E\left[\int_i \sum_{t=0}^{T-1} \beta^t \left(v'(c_{i,t}^b) \cdot D_{i,t}\right) di\right] + \lambda \cdot \frac{\partial G(b)}{\partial b} = 0 \tag{A.29}$$

where the budget effect is

$$\begin{aligned} \frac{\partial G(b)}{\partial b} &= -\int_i E\left[\sum_{t=0}^{T-1} (1+r_t)^{-t} \left(D_{i,t} + \frac{\partial D_{i,t}}{\partial b} \cdot (b+\tau)\right)\right] di && \text{(A.30)} \\ &= -DI\left[1 + \frac{\partial DI}{\partial b} \frac{1}{DI} (b+\tau)\right] && \text{(A.31)} \\ &= -DI\left[1 + \frac{\partial DI}{\partial b} \frac{b}{DI} \frac{1}{1-DI}\right] && \text{(A.32)} \end{aligned}$$

and $DI \equiv \int_i E\left[\sum_{t=0}^{T-1} (1+r_t)^{-t} \left(D_{i,t}\right)\right] di$ is the average DI recipient share in the population.[31] We assume that the planner's budget constraint is differentiable, i.e., $\frac{\partial G(b)}{\partial b}$ exists.

Proof. of (A.29).

[30]The more general problem, where the planner chooses the optimal benefit function $b_s(X_{i,s,})$ in each period s and each $X_{i,s}$, poses no significant additional theoretical challenges. Similar formulas apply. The only difference is that the incidence of the benefit changes is then no longer uniform across all DI recipients. For notational simplicity, we abstract from time-varying benefits since it is irrelevant in our empirical implementation with a uniform DI benefit change.

[31]Equation (A.32) uses the balanced budget constraint, implying $\tau = \frac{DI}{1-DI} \cdot b$.

The first order condition in (A.29) follows from applying Clausen and Strub (2020)'s Differentiable Sandwich Lemma. Clausen and Strub (2020) show that if a function F(b) is sandwiched at some point b between two differentiable functions (upper and lower support functions U(b) and L(b)), then this function F is differentiable at this point b and the derivative of the sandwiched function F equals the derivative of the upper and lower support functions at this point, i.e., $F'(b) = U'(b) = L'(b)$. Suppose b^* is the DI benefit level that maximizes welfare. By definition $W(b^*) \geq W(b)\ \forall b$. We take the constant function $U(b) = W(b^*)$ as the upper support function with $U'(b) = 0$. For the lower support function, we take $L(b) = \int \bar{V}_i(b)di + \lambda\,(G(b))$ where $\bar{V}_i(b)$ denotes the agent's indirect utility if she sticks to her behavior that is optimal for benefits b^* even when benefit levels change. Thus, $\bar{V}_i(b) \leq V_i(b)$ and $L(b)$ is a lower support function for $W(b)$, which implies that

$$\frac{\partial W(b)}{\partial b} = \frac{\partial L(b)}{\partial b} \tag{A.33}$$

$$= E\left[\int_i \sum_{t=0}^{T-1} \beta^t \mu_{i,t}^D \cdot D_{i,t} di\right] + \lambda \cdot \frac{\partial G(b)}{\partial b} \tag{A.34}$$

$$= E\left[\int_i \sum_{t=0}^{T-1} \beta^t \left(v'(c_{i,t}^b) \cdot D_{i,t}\right) di\right] + \lambda \cdot \frac{\partial G(b)}{\partial b} \tag{A.35}$$

$$= \frac{\partial U(b)}{\partial b} = 0 \tag{A.36}$$

and concludes the proof of (A.29). □

If we assume again that $\lambda = \mathbb{E}\left[u'(c^w)|DI = 0\right] \equiv \int_i E\left[\sum_{t=0}^{T-1} (1+r_t)^{-t} \left(1 - D_{i,t}\right) \cdot u'(c_{i,t}^w)\right] di / \left[\int_i E\left[\sum_{t=0}^{T-1} (1+r_t)^{-t} \left(1 - D_{i,t}\right)\right] di\right]$, we arrive from (A.29) at the Baily-Chetty formula

$$\frac{\mathbb{E}\left[v'(c^b)|DI = 1\right]}{\mathbb{E}\left[u'(c^w)|DI = 0\right]} = 1 + \frac{\varepsilon_{DI,b}}{1 - DI}, \tag{A.37}$$

where $\mathbb{E}\left[v'(c^b)|DI = 1\right] \equiv \frac{\int_i E\left[\sum_{t=0}^{T-1} \beta^t \cdot v'(c_{i,t}^b) \cdot D_{i,t}\right] di}{\int_i E\left[\sum_{t=0}^{T-1} (1+r_t)^{-t} \left(D_{i,t}\right)\right] di}$ denotes the average marginal utility of consumption of DI recipients.

Application Decision. In the dynamic setting, the decision to apply for DI benefits becomes forward-looking. At time t the value function of being employed is

$$V_{i,t}^{E}\left(X_{i,t}\right)=u(c_{i,t}^{w})-\psi(\theta_{i,t})+\beta\cdot E\left[\alpha_{i,t}p_{i,t}V_{i,t+1}^{D}+\left(1-\alpha_{i,t}p_{i,t}\right)V_{i,t+1}^{E}|X_{i,t}\right]. \quad \text{(A.38)}$$

Individuals decide in period t whether to apply for DI benefits, $\alpha_{i,t}=1$, or not apply, $\alpha_{i,t}=0$. An applicant is awarded benefits with probability $p_{i,t}$. In case of an award, individuals receive DI benefits in the next period. Hence, individuals employed in period t enter the DI system with probability $\alpha_{i,t}p_{i,t}$ in period $t+1$ and remain employed with probability $1-\alpha_{i,t}p_{i,t}$. The value of being on DI, which is an absorbing state, is given by

$$V_{i,t}^{D}\left(X_{i,t}\right)=v(c_{i,t}^{b})+\beta\cdot E\left[V_{i,t+1}^{D}|X_{i,t}\right]. \quad \text{(A.39)}$$

In period t individuals apply for DI benefits, $\alpha_{i,t}=1$, if the expected value of receiving DI benefits exceeds the expected value of working:

$$\Gamma(\theta_{i,t},A_{i,t},\chi_{i,t})\equiv E\left[V_{i,t+1}^{D}-V_{i,t+1}^{E}|\theta_{i,t},A_{i,t},\chi_{i,t}\right]\geq 0. \quad \text{(A.40)}$$

For given state variables A_t and χ_t, the marginal entrant θ_t^A in period t is defined by

$$\Gamma(\theta_t^A,A_t,\chi_t)=0. \quad \text{(A.41)}$$

$\Gamma(\theta_t^A,A_t,\chi_t)=0$ defines a unique marginal entrant for given A_t and χ_t if the disability shock is persistent. By persistence we mean that a disability shock today reduces the expected value of working, i.e., $\frac{\partial E\left[V_{i,t+1}^{E}|X_{i,t}\right]}{\partial\theta_t}<0$.[32] In this case, θ_t^A conditional

[32]Persistence in disability shocks makes intuitive sense. The alternatives to persistence are harder to motivate. If a disability shock today does not change the expected value of working, $\frac{\partial E\left[V_{i,t+1}^{E}\right]}{\partial\theta_t}=0$, then everybody would always apply, or everybody would never apply (depending on whether working is more attractive than receiving DI benefits, on average). The application decision would thus be independent of the disability level. If a disability shock increases the value of working $\frac{\partial E\left[V_{i,t+1}^{E}\right]}{\partial\theta_t}>0$, then the application decision would be positively correlated with health, i.e., positive health shocks today would increase DI applications and negative health shocks today would decrease the application probability.

on A_t and χ_t is uniquely defined by (A.41), because $E\left[V^D_{i,t+1}|X_{i,t}\right]$ does not depend on the disability level itself, i.e., $\frac{\partial E\left[V^D_{i,t+1}|X_{i,t}\right]}{\partial \theta_t} = 0$, while $\frac{\partial E\left[V^E_{i,t+1}|X_{i,t}\right]}{\partial \theta_t} < 0$ and hence $E\left[V^D_{i,t+1}|X_{i,t}\right]$ and $E\left[V^E_{i,t+1}|X_{i,t}\right]$ have single crossing.

Inflow Responses Reveal Insurance Value of DI. Equation (A.41) implicitly defines the marginal entrant at age t conditional on the state variables A_t and χ_t. We assume $\frac{\partial E\left[V^E_{i,t+1}|X_{i,t}\right]}{\partial \theta_t} < 0$, i.e., the value of continuing working varies smoothly with the disability level ($E\left[V^E_{i,t+1}|X_{i,t}\right]$ is differentiable in θ_t) and a disability shock today reduces the expected value of working. θ^A_t is then differentiable in b by the implicit function theorem. A change in the DI benefit b has the following DI inflow effect at age $t+1$:

$$\frac{\partial}{\partial b} DI_{t+1} = \frac{\partial}{\partial b} \int E\left[\int_{\theta^A_t}^{\infty} p(\theta) dF(\theta) | A_t, \chi_t\right] dG(A_t, \chi_t) \tag{A.42}$$

$$= \int E\left[-\frac{\partial \theta^A_t}{\partial b} p(\theta^A_t) f(\theta^A_t) | A_t, \chi_t\right] dG(A_t, \chi_t), \tag{A.43}$$

where $G(A_t, \chi_t)$ denotes the distribution of the state variables A_t and χ_t. Note that there can be multiple marginal entrants so that different disability cutoffs θ^A_t apply for different state variables A_t and χ_t. Equation (A.43) sums up the responses from all marginal entrants. Furthermore, we take the distribution $G(A_t, \chi_t)$ as given, which means that equation (A.43) quantifies the DI inflow effect at age $t+1$ when individuals learn at age t that DI benefits permanently change. This assumption reflects our empirical approach that sums up inflow effects at different ages from different cohorts in the short run. Equation (A.41), the implicit function theorem, and the envelope theorem imply that

$$\frac{\partial \theta^A_t}{\partial b} = -\frac{\frac{\partial E\left[V^D_{i,t+1} - V^E_{i,t+1} | \theta^A_t, A_t, \chi_t\right]}{\partial b}}{\frac{\partial \Gamma_t}{\partial \theta^A_t}} \tag{A.44}$$

$$= -\frac{1}{\frac{\partial \Gamma_t}{\partial \theta^A_t}} \; E\left[v'(c^b_{t+1}) + \sum_{k=1}^{T-t-2} \beta^k \left(1 - DI_{t+1,t+k}\right) v'(c^b_{t+1+k}) | \theta^A_t, A_t, \chi_t\right] \tag{A.45}$$

where $1 - DI_{f,s} \equiv \prod_{j=f}^{s} \left(1 - \alpha_j p_j\right)$. Combining (A.43) and (A.45) yields

$$\begin{aligned} \frac{\partial}{\partial b} DI_{t+1} &= \int E\left[h(\theta_t^A)\left(v'(c_{t+1}^b) + \sum_{k=1}^{T-t-2} \beta^k \left(1 - DI_{t+1,t+k}\right) v'(c_{t+1+k}^b)\right) |\theta_t^A, A_t, \chi_t\right] dM(\theta_t^A, A_t, \chi_t) \\ &\equiv E\left[v'(c_{t+1}^b) + \sum_{k=1}^{T-t-2} \delta_{t+1}^k v'(c_{t+1+k}^b) |\theta_t^A \in Q^A\right], \end{aligned} \tag{A.46}$$

where $h(\theta_t^A) \equiv \frac{p(\theta_t^A) f(\theta_t^A)}{\partial \Gamma_t / \partial \theta_t^A}$, and $\delta_{t+1}^k \equiv \beta^k \left(1 - DI_{t+1,t+k}\right)$. Therefore, the inflow effect at age $t+1$ corresponds to the ($h(\theta_t^A)$-weighted) average marginal utility of consumption of all marginal entrants at age t when on DI benefits, which we denote by expression (A.46).

Analogously, a permanent change in wages implies

$$\frac{\partial \theta_t^A}{\partial w} = -\frac{\frac{\partial E\left[V_{i,t+1}^D - V_{i,t+1}^E |\theta_t^A, A_t, \chi_t\right]}{\partial w}}{\frac{\partial \Gamma_t}{\partial \theta_t^A}} \tag{A.47}$$

$$= \frac{1}{\frac{\partial \Gamma_t}{\partial \theta_t^A}} E\left[u'(c_{t+1}^w) + \sum_{k=1}^{T-t-2} \beta^k \left(1 - DI_{t+1,t+k}\right) u'(c_{t+1+k}^w) |\theta_t^A, A_t, \chi_t\right] \tag{A.48}$$

and the inflow effect at age $t+1$ is given by

$$\begin{aligned} \frac{\partial}{\partial w} DI_{t+1} &= -\int E\left[h(\theta_t^A)\left(u'(c_{t+1}^w) + \sum_{k=1}^{T-t-2} \beta^k \left(1 - DI_{t+1,t+k}\right) u'(c_{t+1+k}^w)\right) |\theta_t^A, A_t, \chi_t\right] dM(\theta_t^A, A_t, \chi_t) \\ &\equiv -E\left[u'(c_{t+1}^w) + \sum_{k=1}^{T-t-2} \delta_{t+1}^k u'(c_{t+1+k}^w) |\theta_t^A \in Q^A\right]. \end{aligned} \tag{A.49}$$

Our empirical estimate of the DI inflow effect sums up the inflow effect at all ages, i.e.,

$$\frac{\partial DI / \partial b}{-\partial DI / \partial w} = \frac{\sum_{t=0}^{T-2} \frac{\partial}{\partial b} DI_{t+1}}{-\sum_{t=0}^{T-2} \frac{\partial}{\partial w} DI_{t+1}} \tag{A.50}$$

$$= \frac{\sum_{t=0}^{T-2} E\left[v'(c_{t+1}^b) + \sum_{k=1}^{T-t-2} \delta_{t+1}^k v'(c_{t+1+k}^b) |\theta_t^A \in Q^A\right]}{\sum_{t=0}^{T-2} E\left[u'(c_{t+1}^w) + \sum_{k=1}^{T-t-2} \delta_{t+1}^k u'(c_{t+1+k}^w) |\theta_t^A \in Q^A\right]} \tag{A.51}$$

Equation (A.51) indicates that the ratio of DI inflow responses with respect to changes in benefits and wages corresponds to the marginal entrants' ratio of marginal utilities of consumption when on DI benefits versus when working. The ratio of relative inflow responses provides a lower bound for the insurance value of DI benefits if marginal entrants' insurance value is lower in all periods than that of all DI recipients (as we assume in the static model).[33]

Temporary Changes in Wages. We now consider the DI inflow response to temporary wage changes. For a temporary wage change $w_{T,t} = \{dw_{i,k}\}_{k=t}^{T-1}$, where the intensity of the shock $dw_{i,k}$ can vary over time and disappear after some time ($dw_{i,t>s} = 0$), we have

$$\begin{aligned}\frac{\partial\theta_t^A}{\partial w_{T,t+1}} &= -\frac{\frac{\partial E\left[V_{i,t+1}^D - V_{i,t+1}^E | \theta_t^A, A_t, \chi_t\right]}{\partial w_{T,t+1}}}{\frac{\partial\Gamma_t}{\partial\theta_t^A}} \\ &\approx \frac{1}{\frac{\partial\Gamma_t}{\partial\theta_t^A}} E\left[u'(c_{t+1}^w)dw_{i,t+1} + \sum_{k=1}^{T-t-2}\beta^k\left(1 - DI_{t+1,t+k}\right)u'(c_{t+1+k}^w)dw_{i,t+1+k}|\theta_t^A, A_t, \chi_t\right].\end{aligned} \tag{A.52}$$

To make the response of a temporary wage shocks comparable to a permanent benefit change, we rescale the response by the present value of the wage shock $PVW_{t+1} = dw_{i,t+1} + \sum_{k=1}^{T-t-2}\beta^k\left(1 - DI_{t+1,t+k}\right)dw_{i,t+1+k}$. This formulation corresponds to our empirical approach, where we rescale the effects of local labor market shocks by their impact on lifetime income. The question is how this rescaled effect, $\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T}$, compares to the effect of a permanent wage shock, $\frac{\partial\theta_t^A}{\partial w}$. Both $\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T}$ and $\frac{\partial\theta_t^A}{\partial w}$ measure weighted average marginal utilities of consumption of the marginal entrants. The difference is that $\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T}$ puts more weight on periods in the nearer future when the temporary wage shocks are present.

We have $-\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T} \approx -\frac{\partial\theta_t^A}{\partial w}$ if marginal utility of consumption while working is roughly constant across periods for the marginal entrants. When marginal utility of

[33] We can use the mediant theorem to show this statement formally. Equation (A.51) is a sum of the form $\frac{\sum v_j}{\sum u_j}$, where $v_j \equiv h(\theta_{t(j)}^A)\delta_{t(j)+1}^k v'(c_{t(j)+1+k(j)}^b)$ and $u_j \equiv h(\theta_{t(j)}^A)\delta_{t(j)+1}^{k(j)} u'(c_{t(j)+1+k(j)}^w)$ for all t, k-combinations in (A.51) for all marginal entrants (including $k = 0$, where $v_j \equiv h(\theta_{t(j)}^A)v'(c_{t(j)+1}^b)$ and $u_j \equiv h(\theta_{t(j)}^A)u'(c_{t(j)+1}^w)$). Let $m_j = \frac{v_j}{u_j} = \frac{v'(c_{t(j)}^b)}{u'(c_{t(j)}^w)}$ and the indexing j is such that $m_1 \leq \ldots \leq m_j \leq \ldots \leq m_N$. By the mediant theorem, $\frac{\partial DI/\partial b}{\partial DI/\partial w} \leq m_N$, and so if the marginal entrants' insurance value is always lower than that of all DI recipients, we have $\frac{\partial DI/\partial b}{\partial DI/\partial w} \leq m_N \leq \frac{\mathbb{E}\left[v'(c^b)|DI=1\right]}{\mathbb{E}\left[u'(c^w)|DI=0\right]}$.

consumption is falling with age, we have $-\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T} \geq -\frac{\partial\theta_t^A}{\partial w}$. If $-\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T} \geq -\frac{\partial\theta_t^A}{\partial w}$ holds, then the rescaled inflow response to a temporary wage shock is larger than the inflow response to a permanent wage shock. Following the steps above, it then follows that $\frac{\partial DI/\partial b}{\partial DI/\partial w_T} \leq \frac{\partial DI/\partial b}{\partial DI/\partial w} \leq$ Insurance Value and our empirical implementation provides a lower bound for the insurance value of DI benefits. If $-\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T} < -\frac{\partial\theta_t^A}{\partial w}$ holds, then $\frac{\partial DI/\partial b}{\partial DI/\partial w_T}$ is likely still a lower bound for the insurance value. As long as the weighted marginal utility of consumption of marginal entrants that enters into $\frac{1}{PVW_{t+1}}\frac{\partial\theta_t^A}{\partial w_T}$ is larger than the average marginal utility of consumption in the working population, we have $\frac{\partial DI/\partial b}{\partial DI/\partial w_T} \leq \frac{\mathbb{E}\left[v'(c^b)|DI=1\right]}{\mathbb{E}[u'(c^w)|DI=0]} =$ Insurance Value.

A.4.2 Dynamic Model with Three Labor Market States.

Setup. We now extend the model with three labor market states from Appendix A.3 to a dynamic setting. The same notation, timing, and constraints from the dynamic model in Appendix A.4.1 apply. But there is an additional labor market state, unemployment, and decisions about job search and effort on the job have to be made. The laws of motion of assets in the disability, employment and unemployment state are

$$A_{i,t+1} = (1+r_t)A_{i,t} + b(X_{i,t}) - c^b_{i,t}(X_{i,t}) \tag{A.53}$$

$$A_{i,t+1} = (1+r_t)A_{i,t} + w(X_{i,t}) - \tau_{i,t}(X_{i,t}) - c^w_{i,t}(X_{i,t}) \tag{A.54}$$

$$A_{i,t+1} = (1+r_t)A_{i,t} + z(X_{i,t}) - c^z_{i,t}(X_{i,t}) \tag{A.55}$$

where $z(X_{i,t})$ denote unemployment benefits.

The sequence of events and choices is the following. At the beginning of the period, the shocks $\theta_{i,t}$ and $\chi_{i,t}$ are revealed. Having learned $X_{i,t}$, which also includes the labor market state (employed, unemployed, on disability benefits), the individual decides whether to apply to DI, $\alpha_{i,t}(X_{i,t}) = 1$, or not, $\alpha_{i,t}(X_{i,t}) = 0$. If the application is rejected, individuals stay in the labor market and can apply again the next period. Moreover, individuals also decide on how hard to search for a job if unemployed and how much effort to exert on the job. With effort $e_{i,t}$ unemployed individuals are employed next period with probability $q(e_{i,t};\Omega_t)$ that depends both on their search effort and the exogenous economic conditions Ω_t. Employed individuals remain in their job in the next period with probability $s(e_{i,t};\Omega_t)$, which again depends on on-the-job effort $e_{i,t}$ and economic conditions Ω_t.

Denote by $DI_{i,t+1}$ the probability that in period $t+1$ agent i is a DI benefit recipient, which is

$$D_{i,t+1} = 1 - \left[\prod_{k=0}^{t} \left(1 - \alpha_{i,k}(X_{i,k}) \cdot p\left(\theta_{i,k}\right)\right) \right]. \qquad \text{(A.56)}$$

We denote by $E_{i,t+1}$ the probability that agent i is employed in period $t+1$, which is

$$E_{i,t+1} = \left(1 - D_{i,t+1}\right)\left(s(e_{i,t};\Omega_t) + q(e_{i,t};\Omega_t)\right). \qquad \text{(A.57)}$$

The probability to be unemployed is $U_{i,t+1} = 1 - D_{i,t+1} - E_{i,t+1}$. The agent's problem is given by

$$\begin{aligned} V_i(b) &= \max E\left[\sum_{t=0}^{T-1} \beta^t \left(v(c^b_{i,t}) \cdot D_{i,t} + \left(u(c^w_{i,t}) - \psi(\theta_{i,t}, e_{i,t})\right) \cdot E_{i,t} + \left(u(c^z_{i,t}) - \phi(\theta_{i,t}, e_{i,t})\right) \cdot U_{i,t}\right)\right] \\ &+ E\left[\sum_{t=0}^{T-1} \beta^t \mu^D_{i,t} \left((1+r_t)A_{i,t} + b_{i,t} - c^b_{i,t} - A_{i,t+1}\right) D_{i,t}\right] \qquad \text{(A.58)} \\ &+ E\left[\sum_{t=0}^{T-1} \beta^t \mu^W_{i,t} \left((1+r_t)A_{i,t} + w_{i,t} - \tau_{i,t} - c^w_{i,t} - A_{i,t+1}\right) E_{i,t}\right] \\ &+ E\left[\sum_{t=0}^{T-1} \beta^t \mu^Z_{i,t} \left((1+r_t)A_{i,t} + z_{i,t} - c^z_{i,t} - A_{i,t+1}\right) U_{i,t}\right]. \end{aligned}$$

Planner's Problem. The social planner maximizes social welfare by choosing DI benefits and solves

$$\max_b W(b) = \int_i V_i(b) di + \lambda \cdot G(b) \qquad \text{(A.59)}$$

where

$$G(b) = \int_i E\left[\sum_{t=0}^{T-1} (1+r_t)^{-t} \left(E_{i,t} \cdot \tau_{i,t} - D_{i,t} \cdot b_{i,t} - U_{i,t} \cdot z_{i,t}\right)\right] di \qquad \text{(A.60)}$$

is the planners net revenue and λ denotes the Lagrange multiplier on the planner's budget constraint. The solution to the planner's problem is still governed by (A.29). The only difference compared to the two labor market state model is that potential spillover effects to unemployment benefits enter the fiscal effect $\frac{\partial G(b)}{\partial b}$.

Application Decision. The application decision is forward-looking and individuals apply in period t, $\alpha_{i,t} = 1$, if

$$\Gamma(\theta_{i,t}, A_{i,t}, \chi_{i,t}) \equiv E\left[V^D_{i,t+1} - \left(\kappa_{i,t} V^E_{i,t+1} + (1-\kappa_{i,t}) V^U_{i,t+1}\right) | \theta_{i,t}, A_{i,t}, \chi_{i,t}\right] \geq 0 \quad \text{(A.61)}$$

where $\kappa_{i,t} = s_{i,t}$ if the individual is employed in period t and $\kappa_{i,t} = q_{i,t}$ if the individual is unemployed. At time t the value function of being employed is

$$\begin{aligned} V^E_{i,t}\left(X_{i,t}\right) &= u(c^w_{i,t}) - \psi(\theta_{i,t}, e_{i,t}) \ldots && \text{(A.62)} \\ &\quad +\beta \cdot E\left[\alpha_{i,t} p_{i,t} V^D_{i,t+1} + \left(1 - \alpha_{i,t} p_{i,t}\right)\left[s_{i,t} V^E_{i,t+1} + \left(1 - s_{i,t}\right) V^U_{i,t+1}\right] | X_{i,t}\right], \end{aligned}$$

where $s_{i,t} \equiv s(e_{i,t}; \Omega_t)$. The value function of being unemployed is

$$\begin{aligned} V^U_{i,t}\left(X_{i,t}\right) &= u(c^z_{i,t}) - \phi(\theta_{i,t}, e_{i,t}) \ldots && \text{(A.63)} \\ &\quad +\beta \cdot E\left[\alpha_{i,t} p_{i,t} V^D_{i,t+1} + \left(1 - \alpha_{i,t} p_{i,t}\right)\left[q_{i,t} V^E_{i,t+1} + \left(1 - q_{i,t}\right) V^U_{i,t+1}\right] | X_{i,t}\right], \end{aligned}$$

where $q_{i,t} \equiv q(e_{i,t}; \Omega_t)$. The value of being on DI, which is an absorbing state, is given by

$$V^D_{i,t}\left(X_{i,t}\right) = v(c^b_{i,t}) + \beta \cdot E\left[V^D_{i,t+1} | X_{i,t}\right]. \quad \text{(A.64)}$$

For given state variables A_t and χ_t, the marginal entrant θ^A_t in period t is defined by

$$\Gamma(\theta^A_t, A_t, \chi_t) = 0. \quad \text{(A.65)}$$

$\Gamma(\theta^A_t, A_t, \chi_t) = 0$ defines a unique marginal entrant for given A_t and χ_t if a disability shock today reduces the expected value of staying active in the labor market, i.e., $\frac{\partial E\left[\kappa_{i,t} V^E_{i,t+1} + (1-\kappa_{i,t}) V^U_{i,t+1} | X_{i,t}\right]}{\partial \theta_t} < 0$. In this case θ^A_t conditional on A_t and χ_t is uniquely defined by (A.65) because $E\left[V^D_{i,t+1} | X_{i,t}\right]$ does not depend on the disability level itself, i.e., $\frac{\partial E\left[V^D_{i,t+1} | X_{i,t}\right]}{\partial \theta_t} = 0$, while $\frac{\partial E\left[\kappa_{i,t} V^E_{i,t+1} + (1-\kappa_{i,t}) V^U_{i,t+1} | X_{i,t}\right]}{\partial \theta_t} < 0$ and hence $E\left[V^D_{i,t+1} | X_{i,t}\right]$ and $E\left[\kappa_{i,t} V^E_{i,t+1} + (1-\kappa_{i,t}) V^U_{i,t+1} | X_{i,t}\right]$ have single crossing.

A change in the DI benefit b has the following DI inflow effect at age $t+1$:

$$\frac{\partial}{\partial b} DI_{t+1} = \frac{\partial}{\partial b} \int E\left[\int_{\theta_t^A}^{\infty} p(\theta) dF(\theta) | A_t, \chi_t\right] dG(A_t, \chi_t) \tag{A.66}$$

$$= \int E\left[-\frac{\partial \theta_t^A}{\partial b} p(\theta_t^A) f(\theta_t^A) | A_t, \chi_t\right] dG(A_t, \chi_t), \tag{A.67}$$

where $G(A_t, \chi_t)$ denotes the distribution of the state variables A_t and χ_t. Equation (A.67) sums up the responses from all marginal entrants. As in Appendix, A.4.1 we take the distribution $G(A_t, \chi_t)$ as given. Equation (A.65), the implicit function theorem, and the envelope theorem imply that

$$\frac{\partial \theta_t^A}{\partial b} = -\frac{\frac{\partial E\left[V_{i,t+1}^D - \left(\kappa_{i,t} V_{i,t+1}^E + (1-\kappa_{i,t}) V_{i,t+1}^U\right) | \theta_{i,t}, A_{i,t}, \chi_{i,t}\right]}{\partial b}}{\frac{\partial \Gamma_t}{\partial \theta_t^A}} \tag{A.68}$$

$$= -\frac{1}{\frac{\partial \Gamma_t}{\partial \theta_t^A}} E\left[v'(c_{t+1}^b) + \sum_{k=1}^{T-t-2} \beta^k \left(1 - DI_{t+1,t+k}\right) v'(c_{t+1+k}^b) | \theta_t^A, A_t, \chi_t\right], \tag{A.69}$$

where $1 - DI_{f,s} \equiv \prod_{j=f}^{s} \left(1 - \alpha_j p_j\right)$. Combining (A.67) and (A.69) yields

$$\frac{\partial}{\partial b} DI_{t+1} = \int E\left[h(\theta_t^A)\left(v'(c_{t+1}^b) + \sum_{k=1}^{T-t-2} \beta^k \left(1 - DI_{t+1,t+k}\right) v'(c_{t+1+k}^b)\right) | \theta_t^A, A_t, \chi_t\right] dM(\theta_t^A, A_t, \chi_t)$$

$$\equiv E\left[v'(c_{t+1}^b) + \sum_{k=1}^{T-t-2} \delta_{t+1}^k v'(c_{t+1+k}^b) | \theta_t^A \in Q^A\right], \tag{A.70}$$

where $h(\theta_t^A) \equiv \frac{p(\theta_t^A) f(\theta_t^A)}{\partial \Gamma_t / \partial \theta_t^A}$ and $\delta_{t+1}^k \equiv \beta^k \left(1 - DI_{t+1,t+k}\right)$. Therefore, as in Appendix A.4.1 the inflow effect at age $t+1$ corresponds to the (weighted) average marginal utility of consumption of all marginal entrants at age t when on DI benefits, which we denote by expression (A.70).

Analogously, a permanent change in wages implies

$$\frac{\partial \theta_t^A}{\partial w} = -\frac{\frac{\partial E\left[V_{i,t+1}^D - \left(\kappa_{i,t} V_{i,t+1}^E + (1-\kappa_{i,t}) V_{i,t+1}^U\right) | \theta_{i,t}, A_{i,t}, \chi_{i,t}\right]}{\partial w}}{\frac{\partial \Gamma_t}{\partial \theta_t^A}} \tag{A.71}$$

$$= \frac{1}{\frac{\partial \Gamma_t}{\partial \theta_t^A}} E\left[E_t u'(c_{t+1}^w) + \sum_{k=1}^{T-t-2} \beta^k E_{t+k} u'(c_{t+1+k}^w) | \theta_t^A, A_t, \chi_t\right], \tag{A.72}$$

where E_t denotes the share of employed in period $t+1$. The inflow effect at age $t+1$ is given by

$$\begin{aligned}\frac{\partial}{\partial w}DI_{t+1} &= -\int E\left[h(\theta_t^A)\left(E_t u'(c_{t+1}^w)+\sum_{k=1}^{T-t-2}\beta^k E_{t+k}u'(c_{t+1+k}^w)\right)|\theta_t^A, A_t, \chi_t\right]dM(\theta_t^A, A_t, \chi_t) \\ &\equiv -E\left[E_t u'(c_{t+1}^w)+\sum_{k=1}^{T-t-2}\omega_{t+1}^k u'(c_{t+1+k}^w)|\theta_t^A\in Q^A\right] \qquad \text{(A.73)}\end{aligned}$$

where $\omega_{t+1}^k \equiv \beta^k E_{t+k}$. Our empirical estimate of the DI inflow effect sums up the inflow effect at all ages, i.e.,

$$\begin{aligned}\frac{\partial DI/\partial b}{-\partial DI/\partial w} &= \frac{\sum_{t=0}^{T-2}\frac{\partial}{\partial b}DI_{t+1}}{-\sum_{t=0}^{T-2}\frac{\partial}{\partial w}DI_{t+1}} \qquad \text{(A.74)} \\ &= \frac{\sum_{t=0}^{T-2}E\left[v'(c_{t+1}^b)+\sum_{k=1}^{T-t-2}\delta_{t+1}^k v'(c_{t+1+k}^b)|\theta_t^A\in Q^A\right]}{\sum_{t=0}^{T-2}E\left[E_t u'(c_{t+1}^w)+\sum_{k=1}^{T-t-2}\omega_{t+1}^k u'(c_{t+1+k}^w)|\theta_t^A\in Q^A\right]}. \qquad \text{(A.75)}\end{aligned}$$

As in Appendix A.4.1, the ratio of inflow responses with respect to benefits and wages is a ratio of the weighted average marginal utility of consumption when on DI and the weighted average marginal utility of consumption when working for marginal entrants. We again assume that the insurance value of marginal entrants provides a lower bound for the insurance value of DI benefits in the population.

Temporary Economic Shocks. We now consider temporary economic shocks Ω that affect wages and employment probabilities analogous to the static model from Appendix A.4.2. The temporary economic shocks $\Omega = \{d\Omega_k\}_{k=t}^{T-1}$ change the economic conditions in period k with intensity $d\Omega_k$ (this can also be zero, i.e., no effect in period k). The economic shock affects the wage, $\frac{\partial w_{i,t}}{\partial \Omega}$, the layoff probability, $\frac{\partial(1-s_{i,t})}{\partial \Omega}$, and the job finding probability, $\frac{\partial q_{i,t}}{\partial \Omega}$. These temporary economic shocks Ω capture our empirical design where we use Bartik IVs as well as oil price shocks to leverage regional exposure to temporary economic shocks.

The effect of a temporary economic shock Ω on DI inflow is given by

$$\frac{\partial}{\partial \Omega}DI_{t+1} = \int E\left[-\frac{\partial \theta_t^A}{\partial \Omega}p(\theta_t^A)f(\theta_t^A)|A_t, \chi_t\right]dG(A_t, \chi_t), \qquad \text{(A.76)}$$

where

$$\frac{\partial \theta_t^A}{\partial \Omega} \equiv -\frac{\frac{\partial E\left[V^D_{i,t+1}-\left(\kappa_{i,t}V^E_{i,t+1}+(1-\kappa_{i,t})V^U_{i,t+1}\right)|\theta_{i,t},A_{i,t},\chi_{i,t}\right]}{\partial \Omega}}{\frac{\partial \Gamma_t}{\partial \theta_t^A}} \tag{A.77}$$

$$= \left(\frac{\partial \Gamma_t}{\partial \theta_t^A}\right)^{-1} E\left[\sum_{k=0}^{T-t-2}\beta^k\left[E_{i,t+k}u'(c^E_{i,t+1+k})\frac{\partial w_{t+1+k}}{\partial \Omega}\right]|\theta_t^A,A_t,\chi_t\right]+ \tag{A.78}$$

$$\left(\frac{\partial \Gamma_t}{\partial \theta_t^A}\right)^{-1} E\left[\sum_{k=0}^{T-t-2}\beta^k\left[\frac{\partial E_{i,t+k}}{\partial \Omega}\left[u(c^E_{i,t+1+k})-\psi_{i,t+1+k}-\left(u(c^U_{i,t+1+k})-\phi_{i,t+1+k}\right)\right]\right]|\theta_t^A,A_t,\chi_t\right],$$

and $\psi_{i,t} \equiv \psi(\theta_{i,t},e_{i,t})$ and $\phi_{i,t} \equiv \phi(\theta_{i,t},e_{i,t})$. $E_{i,t}$ denotes the probability to start out employed in period $t+1$ and $\frac{\partial E_{i,t+k}}{\partial \Omega}$ denotes the change in the employment probability due to the economic shock Ω.[34] If we bound the flow utility of job loss analogously to the static model in Appendix A.3 , $u(c^E_{i,t+k})-\psi_{i,t+k}-\left(u(c^U_{i,t+k})-\phi_{i,t+k}\right) \geq u'(c^E_{i,t+k})\left(w_{i,t+k}-z_{i,t+k}\right)$, we have

$$\frac{1}{PVW_t}\frac{\partial \theta_t^A}{\partial \Omega} \geq \left(\frac{\partial \Gamma_t}{\partial \theta_t^A}\right)^{-1}\frac{1}{PVW_t}E\left[\sum_{k=0}^{T-t-2}\beta^k\left[u'(c^E_{i,t+1+k})\left(E_{i,t+k}\frac{\partial w_{t+1+k}}{\partial \Omega}+\frac{\partial E_{i,t+k}}{\partial \Omega}\left(w_{i,t+1+k}-z_{i,t+1+k}\right)\right)\right]|\theta_t^A,A_t,\chi_t\right],$$

where $PVW_t = E\left[\sum_{k=0}^{T-t-2}\beta^k\left[\left(E_{i,t+k}\frac{\partial w_{t+1+k}}{\partial \Omega}+\frac{\partial E_{i,t+k}}{\partial \Omega}\left(w_{i,t+1+k}-z_{i,t+1+k}\right)\right)\right]|\theta_t^A,A_t,\chi_t\right]$ is the present value effect of the economic shock on income. The ratio of the inflow response with respect to benefits and economics shocks rescaled by their present value impact identifies the ratio of marginal utilities of the marginal entrants

$$\frac{\partial DI/\partial b}{-\partial DI/\partial \Omega} = \frac{\frac{1}{PVB}\sum_{t=0}^{T-2}\frac{\partial}{\partial b}DI_{t+1}}{\frac{1}{PVW}\sum_{t=0}^{T-2}\frac{\partial}{\partial \Omega}DI_{t+1}} \tag{A.79}$$

$$\leq \frac{\sum_{t=0}^{T-2}\frac{1}{PVB_t}E\left[v'(c^b_{t+1})+\sum_{k=1}^{T-t-2}\delta^k_{t+1}v'(c^b_{t+1+k})|\theta_t^A\in Q^A\right]}{\sum_{t=0}^{T-2}\frac{1}{PVW_t}E\left[\sum_{k=0}^{T-t-2}\beta^k\left[u'(c^E_{i,t+1+k})\left(E_{i,t+k}\frac{\partial w_{t+1+k}}{\partial \Omega}+\frac{\partial E_{i,t+k}}{\partial \Omega}\left(w_{i,t+1+k}-z_{i,t+1+k}\right)\right)\right]|\theta_t^A\in Q^A\right]}$$

$$\equiv \frac{\mathbb{E}\left[v'(c^b)|\theta_t^A\in Q^A\right]}{\mathbb{E}\left[u'(c^w)|\theta_t^A\in Q^A\right]}. \tag{A.80}$$

This is again a lower bound for the insurance value if the marginal entrant values DI benefits less than the average DI recipient.

[34] For illustration take an individual who is employed in period t. The probability to be employed in period $t+1$ is given by $E_{i,t} = s_{i,t}$, the probability to be employed in $t+2$ is $E_{i,t+1} = s_{i,t}\cdot s_{i,t+1}+(1-s_{i,t})\cdot q_{i,t+1}$ and so on. An economic shock that starts in period t has the following effect on employment in $t+2$ for individuals being employed in t: $\frac{\partial E_{i,t+2}}{\partial \Omega} = \frac{\partial s_{i,t}}{\partial \Omega}\cdot(s_{i,t+1}-q_{i,t+1})+s_{i,t}\cdot\frac{\partial s_{i,+1}}{\partial \Omega}+(1-s_{i,t})\cdot\frac{\partial q_{i,t+1}}{\partial \Omega}$.

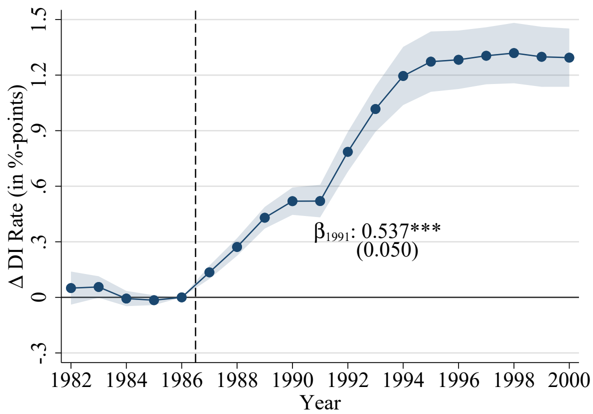
B Additional Results for DI Benefit Generosity

Table B.1: Summary Statistics, Benefit Generosity Sample

	ROC		Quebec	
	1982-1986	1987-1992	1982-1986	1987-1992
A. Outcomes				
DI benefit receipt (in %)	1.25	2.40	1.33	2.02
Non-employment (in %)	15.61	15.84	19.26	18.24
DI benefits of recipients	5,000	9,000	6,700	9,000
	(3,400)	(5,600)	(4,800)	(4,500)
Earnings	16,700	24,600	15,000	21,700
	(18,600)	(34,000)	(15,100)	(24,500)
B. Characteristics				
Share Female	0.51	0.52	0.51	0.52
Share Married	0.66	0.71	0.62	0.63
Age	34.05	39.45	34.11	39.51
	(8.98)	(8.95)	(8.75)	(8.72)
No. Kids	0.28	0.25	0.28	0.22
	(0.65)	(0.59)	(0.66)	(0.56)
No. Observations	6,491,185	7,728,240	2,315,605	2,747,770

Notes: The table reports summary statistics for the Rest of Canada (ROC) and Quebec in the years before (1982-1986) and after the reform (1987-1992). All monetary outcomes are expressed in 2019 dollars using the national CPI. Standard deviations of continuous variables in parentheses. Dollar amounts are rounded to $100 following the vetting guidelines of the research data center.

Figure B.1: Difference in DI Take-Up Between RoC and Quebec by Year excluding Widows



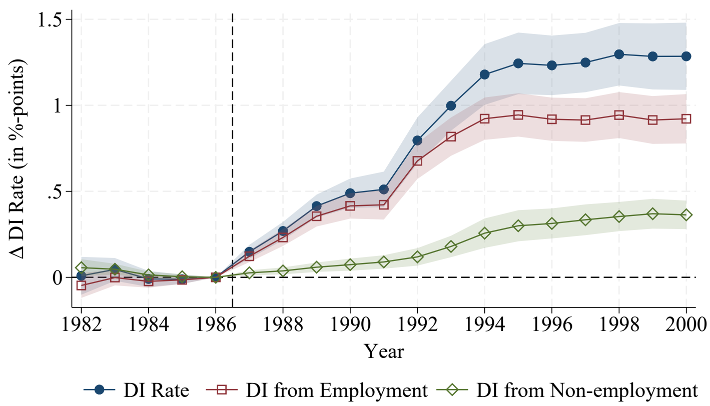
Notes: The figure shows the estimated β_s-coefficients from the econometric specification in (10), excluding individuals whose spouse died in the same year to exclude survivor benefits. The shaded area represents the 95-percent confidence interval.

Table B.2: Impacts of 1987 Reform on DI Receipt, Robustness

	Main estimate	Oil and gas provinces	Unbalanced sample
Coeff. estimate (β_{1991})	0.527***	0.404***	0.534***
	(0.061)	(0.056)	(0.047)
Mean ($\overline{y}$)	1.84	1.59	2.15
Elasticity $\left(\beta_{1991} \cdot \frac{\overline{b}}{\overline{y}}\right)$	0.580***	0.507***	0.398***
	(0.067)	(0.070)	(0.035)
No. Observations	17,543,255	9,232,205	28,433,095

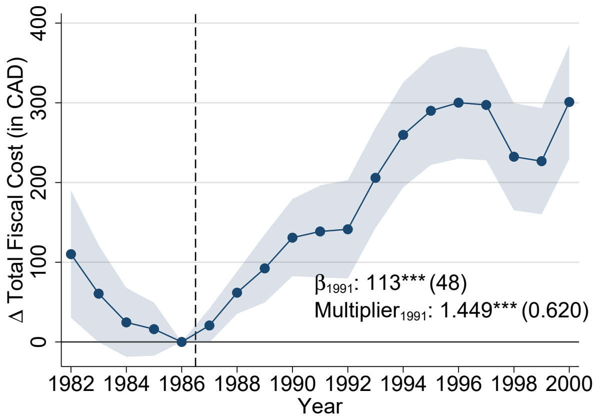
Notes: The table reports the reform effect in the fifth year after the reform for DI receipt (the β_{1991}-coefficient from equation 10). Coefficient estimates and standard errors are multiplied by 100 and should be interpreted as percentage points. The main sample includes 15-59 year-old individuals who filed taxes continuously from 1982 until 1991 or until death, whichever occurred first. Oil and gas provinces sample restricts the main sample to individuals residing in Alberta, British Columbia, Saskatchewan, Newfoundland and Labrador, and Quebec. The unbalanced sample includes all 15-59 year-old individuals from 1982 until 1991, regardless of whether they filed taxes continuously. Standard errors are reported in parentheses and are clustered at the Census Division. All specifications include year fixed effects, province fixed effects, age and its square, and the provincial unemployment rate. Levels of significance: $^{\star}10\%$, $^{\star\star}5\%$, and $^{\star\star\star}1\%$.

Figure B.2: DI receipt from employment and non-employment



Notes: The figure reports the estimated β_s-coefficients from equation (10) for different outcomes."DI Rate" is an indicator equal to one if an individual receives a public disability pension. "DI from Employment" is an indicator equal to one if the individual receives a public disability pension and had positive labor income in the two years before entering DI. "DI from Non-Employment" is an indicator equal to one if the individual receives a public disability pension and had no labor income in the two years before entering DI. The shaded area represents the 95-percent confidence interval.

Figure B.3: Total fiscal cost and fiscal multiplier



Notes: The figure reports the estimated β_s-coefficients from equation (10) for total fiscal cost, defined as DI benefit expenditures minus labor income tax revenues. The shaded area represents the 95-percent confidence interval.

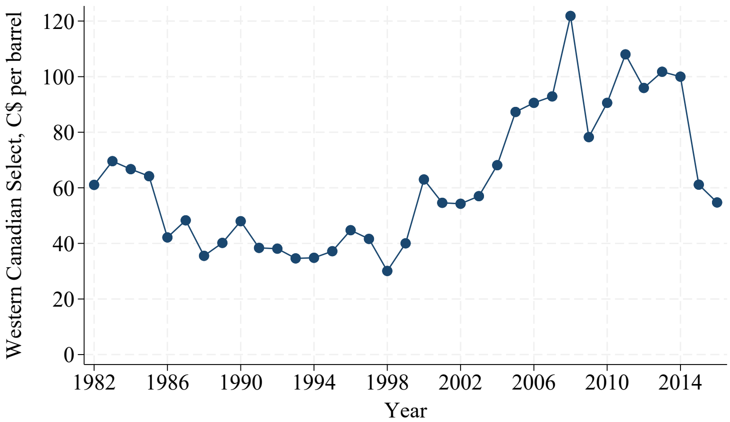
Table B.3: Comparison Marginal vs. Average Entrant, Benefit Change

	Marginal entrant	Average entrant	P-value
Age	50.57***	45.66***	0.000
	(0.81)	(0.13)	
Family size	3.60***	3.04***	0.027
	(0.26)	(0.02)	
Share female	0.426***	0.548***	0.001
	(0.038)	(0.006)	
Pr(died by age 60)	0.132***	0.141***	0.573
	(0.017)	(0.002)	
Share earnings 1. quartile	0.377***	0.290***	0.057
	(0.047)	(0.005)	
Share earnings 2. quartile	0.274***	0.270***	0.854
	(0.026)	(0.005)	
Share earnings 3. quartile	0.213***	0.246***	0.286
	(0.032)	(0.004)	
Share earnings 4. quartile	0.136***	0.193***	0.060
	(0.034)	(0.006)	
Share income 1. quartile	0.357***	0.287***	0.123
	(0.047)	(0.006)	
Share income 2. quartile	0.268***	0.269***	0.988
	(0.030)	(0.005)	
Share income 3. quartile	0.223***	0.243***	0.374
	(0.025)	(0.004)	
Share income 4. quartile	0.152***	0.199***	0.104
	(0.033)	(0.005)	

Notes: The table reports the complier characteristics (β-estimates of equation (13)) for a benefit change and compares the marginal entrant to the average entrant. Earnings and income are measured in $t - 3$. The last column reports the p-value for the null hypothesis that marginal entrant characteristics are identical to characteristics of the average entrant. Levels of significance: *10%, **5%, and ***1%.

C Additional Results for Wage Shocks

Figure C.1: Oil Price over Time



Notes: The figure plots the real oil price trends between 1982 and 2016. We use Western Canadian Select (WCS) prices per barrel, published by the Canadian Association of Petroleum Producers (https://www.capp.ca/en/capp-data-centre/).

Table C.1: Summary Statistics, Wage Change Samples

	1982–2001	2006–2016
A. Outcomes		
DI benefit receipt (in %)	2.71	2.74
Change in DI benefit receipt (in p.p.)	0.94	0.82
Non-employment (in %)	20.13	18.57
Change in non-employment (in p.p.)	1.28	1.31
Avg. lifetime income	36,600	45,600
	(64,300)	(79,500)
Change in avg. lifetime income	3,800	5,500
	(22,700)	(37,400)
DI benefits of recipients	8,300	8,700
	(4,700)	(4,900)
Earnings	39,400	46,900
	(68,700)	(92,300)
B. Characteristics		
Share Female	0.51	0.52
Share Married	0.63	0.62
Age	39.59	41.55
	(10.20)	(10.82)
No. Kids	0.29	0.35
	(0.64)	(0.70)
No. Observations	9,973,095	8,856,110

Notes: The table reports summary statistics for the wage change samples in 1982-2001 and 2006-2016. Standard deviations of continuous variables in parentheses. All monetary outcomes are expressed in 2019 dollars using the national CPI. Dollar amounts are rounded to $100 following the vetting guidelines of the research data center.

Table C.2: OLS, 2SLS, and Reduced-Form Estimates for Different Outcomes

	OLS	2SLS	Reduced-Form
Δ DI enrollment	-0.015***	-0.096***	-5.09***
	(0.003)	(0.018)	(0.82)
Δ DI w/o survivor	-0.014***	-0.095***	-5.01***
	(0.003)	(0.018)	(0.82)
log(DI benefit)	-0.001***	-0.009***	-0.457***
	(0.000)	(0.002)	(0.078)
Pr(moving)	0.003	-0.271	-14.36
	(0.002)	(0.576)	(31.12)
Obs.	18,829,205	18,829,205	18,829,205

Notes: The first two columns report the OLS and 2SLS estimates of equation (11) when we instrument for ΔE_{ict} using the industry shift-share instrument. The last column reports the reduced-form that regresses an outcome variable on the industry shift-share instrument. Coefficient estimates and standard errors for ΔDI enrollment, ΔDI w/o survivor, and Pr(moving) are multiplied by 100 and should be interpreted as percentage points. Standard errors are reported in parentheses and are clustered at the Census Division (CD). The sample includes 15-59 year-old individuals between 1982 and 2016. Levels of significance: *10%, **5%, and ***1%.

Table C.3: The Impact of Earnings/Income on Disability Insurance Enrollment: Oil IV

	Earnings ($1,000)	Current income ($1,000)	Lifetime income ($1,000)
A. Oil employment			
Δ DI	-0.033⋆⋆⋆	-0.034⋆⋆⋆	-0.061⋆⋆⋆
enrollment	(0.009)	(0.009)	(0.015)
1st-stage	212.0⋆⋆⋆	210.0⋆⋆⋆	117.0⋆⋆⋆
coefficient	(17.6)	(17.4)	(9.3)
F-statistic	145.9	145.3	156.3
Obs.	18,829,205	18,829,205	18,829,205
B. Oil price			
Δ DI	-0.022⋆⋆⋆	-0.023⋆⋆⋆	-0.072⋆⋆⋆
enrollment	(0.002)	(0.002)	(0.007)
1st-stage coefficients			
Current	5.0⋆⋆⋆	5.0⋆⋆⋆	1.0⋆⋆⋆
	(0.9)	(0.9)	(0.2)
1st lag	1.0	1.0	1.0⋆⋆⋆
	(1.8)	(1.8)	(0.5)
2nd lag	3.0⋆⋆	3.0⋆⋆	0.0
	(1.3)	(1.3)	(0.2)
F-statistic	22.5	22.5	27.7
Obs.	8,953,440	8,953,440	8,953,440

Notes: The table reports the first-stage and 2SLS estimates of equation (11) for different earnings and income measures when we instrument for ΔE_{ict} using predicted oil employment (panel A) or the oil price (panel B). Coefficient estimates and standard errors for ΔDI enrollment are multiplied by 100 and should be interpreted as percentage points. Standard errors are reported in parentheses and are clustered at the Census Division (CD). The sample includes 15-59 year-old individuals between 1982 and 2016. Levels of significance: ⋆10%, ⋆⋆5%, and ⋆⋆⋆1%.

Table C.4: The Impact of Lifetime Income on Disability Insurance Enrollment, Robustness

	OLS	Industry share	Oil employment	Oil price
A. Years 1982-2001				
Coefficient	-0.016***	-0.082***	-0.084***	-0.052***
enrollment	(0.004)	(0.024)	(0.023)	(0.010)
Effective F-statistic		46.5	14.9	17.4
Obs.	9,973,095	9,973,095	9,973,095	3,386,005
B. Oil and gas provinces				
Coefficient	-0.013***	-0.035**	-0.031**	-0.043***
enrollment	(0.005)	(0.016)	(0.016)	(0.007)
Effective F-statistic		23.0	192.8	34.0
Obs.	9,412,320	9,412,320	9,412,320	4,371,960

Notes: The table reports the OLS and 2SLS estimates of equation (11), instrumenting ΔE_{ict} (lifetime income) using predicted employment based on industry employment shares, oil employment shares, or oil prices. Coefficient estimates and standard errors are multiplied by 100 and should be interpreted as percentage points. Standard errors are reported in parentheses and are clustered at the Census Division (CD). The sample includes 15-59 year-old individuals between 1982 and 2001 (Panel A) and 15-59 year-old individuals between 1982 and 2016 who reside in oil and gas provinces (Alberta, British Columbia, Saskatchewan, and Newfoundland and Labrador) or Quebec (Panel B). Levels of significance: *10%, **5%, and ***1%.

Table C.5: Relation between Industry/Oil Shares and Census Division Characteristics

	Industry shares			Oil shares
	Technical services	Farming	Transportation	
Age 15-19	0.062	-0.350	-0.092	-0.160
	(0.038)	(0.425)	(0.064)	(0.245)
Age 20-24	-0.009	-0.113	-0.055	-0.131
	(0.046)	(0.373)	(0.056)	(0.254)
Age 25-29	0.002	0.181	-0.217***	-0.084
	(0.051)	(0.351)	(0.069)	(0.248)
Age 30-34	-0.107***	-0.195	-0.010	-0.253
	(0.035)	(0.340)	(0.057)	(0.231)
Age 35-39	0.054	-1.559***	0.071	-0.209
	(0.040)	(0.446)	(0.077)	(0.259)
Age 40-44	0.067	0.381	-0.171	-0.228
	(0.096)	(0.814)	(0.138)	(0.236)
Age 45-49	0.003	-1.501*	-0.040	-0.242
	(0.076)	(0.831)	(0.097)	(0.257)
Age 50-54	-0.021	0.283	-0.063	-0.312
	(0.073)	(0.711)	(0.132)	(0.251)
Age 55-59	0.017	1.292**	-0.017	-0.166
	(0.104)	(0.597)	(0.112)	(0.250)
Pr(died by age 60)	0.463	-17.488	-1.597	0.183
	(1.618)	(12.056)	(1.940)	(1.577)
Share earnings 2. quartile	-0.046	0.620	0.039	0.012
	(0.047)	(0.504)	(0.086)	(0.016)
Share earnings 3. quartile	-0.001	1.461**	0.052	1.001
	(0.069)	(0.583)	(0.130)	(0.856)
Share earnings 4. quartile	0.109	-0.680	-0.007	0.581
	(0.100)	(0.604)	(0.117)	(1.228)
Share income 2. quartile	0.065	-0.258	-0.055	0.053
	(0.048)	(0.585)	(0.097)	(0.328)
Share income 3. quartile	0.030	-1.514**	0.007	-0.458
	(0.072)	(0.646)	(0.133)	(1.020)
Share income 4. quartile	-0.038	0.687	0.022	-0.422
	(0.106)	(0.665)	(0.133)	(1.173)
Familiy size	-0.014***	-0.008	-0.001	-0.009**
	(0.003)	(0.024)	(0.004)	(0.004)
Share married	-0.015	0.676***	0.021	0.047***
	(0.018)	(0.133)	(0.030)	(0.017)
Share female	0.095*	-0.376	0.119**	0.001
	(0.055)	(0.319)	(0.058)	(0.045)
R^2	0.89	0.73	0.41	0.25

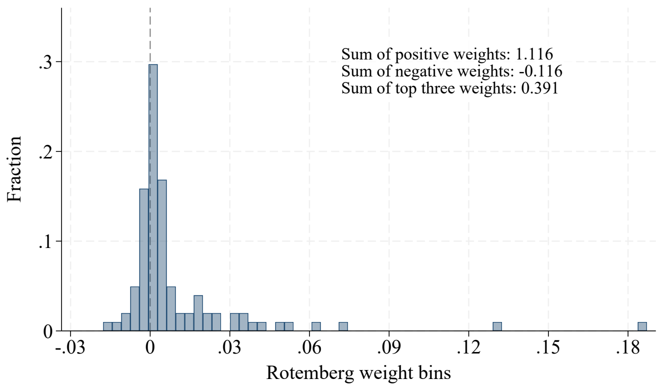
Notes: Each column reports the estimates of a single regression of the initial period the complier characteristics (β-estimates of equation (13)) for benefit and wage changes. Levels of significance: *10%, **5%, and ***1%. Notes: Each column reports results of a single regression of a 1981 industry share on 1982 characteristics. Standard errors in parentheses.

Table C.6: Alternative Estimators and Overidentification Tests

	Coefficient	Over ID test
2SLS (Bartik instrument)	-0.096***	
	(0.018)	
2SLS (many instruments)	-0.088***	1,020.99
	(0.012)	[0.000]
MB2SLS	-0.092***	
	(0.013)	
LIML	-0.114***	2,491.45
	(0.020)	[0.000]

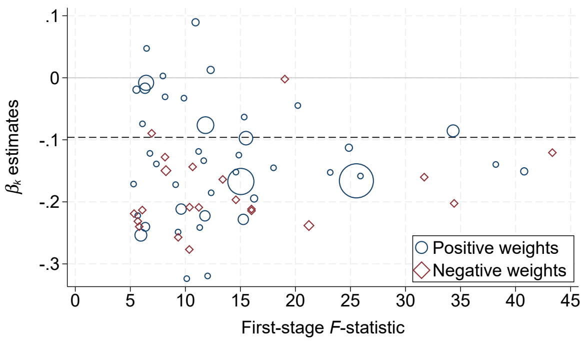
Notes: The table reports the estimates of equation (11) for different estimators, instrumenting for ΔE_{ict} (lifetime income) using predicted employment based on industry employment shares. The dependent variable is the change in DI receipt. The 2SLS estimate with the Bartik instrument reports the main individual-level estimate. The many-instrument 2SLS, MB2SLS, and LIML rows use each industry share (times time period) separately as instruments. Coefficient estimates and standard errors are multiplied by 100 and should be interpreted as percentage points. Standard errors are reported in parentheses and are clustered at the Census Division (CD). P-values for the overidentification tests are reported in brackets. The sample includes 15-59 year-old individuals between 1982 and 2016. Levels of significance: *10%, **5%, and ***1%.

Figure C.2: Distribution of Rotemberg Weights



Notes: This figure shows the full distribution of Rotemberg weights by industry, with weights aggregated across years following Goldsmith-Pinkham et al. (2020).

Figure C.3: Heterogeneity of β_k-estimates



Notes: The figure relates each industry-share instrument's estimated $\hat{\beta}_k$ to its first-stage strength and Rotemberg weight. Each point represents one industry-share instrument, with $\hat{\beta}_k$ plotted on the y-axis and the corresponding first-stage F-statistic on the x-axis. Marker size is proportional to the absolute value of the Rotemberg weight. Circles indicate positive Rotemberg weights, while diamonds indicate negative weights. The horizontal dashed line shows the overall $\hat{\beta}_k$ from the 2SLS (Bartik) estimate reported in column 2 of Table 2. Instruments with first-stage F-statistics below 5 are excluded.

Table C.7: Comparison Marginal vs. Average Entrant, Wage Change (Industry Share Instrument)

	Marginal entrant	Average entrant	P-value
Age	53.12***	45.69***	0.001
	(2.15)	(0.16)	
Family size	3.14***	2.69***	0.011
	(0.17)	(0.02)	
Share female	0.409***	0.585***	0.001
	(0.054)	(0.003)	
Pr(died by age 60)	0.119***	0.092***	0.143
	(0.019)	(0.001)	
Share earnings 1. quartile	0.441***	0.252***	0.001
	(0.058)	(0.003)	
Share earnings 2. quartile	0.256***	0.286***	0.575
	(0.050)	(0.005)	
Share earnings 3. quartile	0.238***	0.262***	0.482
	(0.034)	(0.002)	
Share earnings 4. quartile	0.065	0.200***	0.040
	(0.063)	(0.005)	
Share income 1. quartile	0.345***	0.212***	0.001
	(0.040)	(0.003)	
Share income 2. quartile	0.292***	0.308***	0.770
	(0.053)	(0.003)	
Share income 3. quartile	0.264***	0.271***	0.823
	(0.030)	(0.002)	
Share income 4. quartile	0.099	0.210***	0.086
	(0.062)	(0.004)	

Notes: The table reports the complier characteristics (β-estimates of equation (13)) for wage changes and compares the marginal entrant to the average entrant. Earnings and income are measured in $t-3$. The change in DI receipt on the right-hand side of equation (13) is instrumented with the industry shift share instrument. The last column reports the p-value for the null hypothesis that marginal entrant characteristics are identical to characteristics of the average entrant. Levels of significance: *10%, **5%, and ***1%.

Table C.8: Comparison of Complier Characteristics, Wage Change (Oil Instrument)

	Benefit change		Wage change		Equality test
	Coeff.	Std. err.	Coeff.	Std. err.	(P-value)
Age	50.57***	0.81	49.43***	0.93	0.356
Family size	3.60***	0.26	2.82***	0.09	0.004
Share female	0.426***	0.038	0.545***	0.014	0.003
Pr(died by age 60)	0.132***	0.017	0.078***	0.006	0.003
Share earnings 1. quartile	0.377***	0.047	0.376***	0.020	0.983
Share earnings 2. quartile	0.274***	0.026	0.189***	0.034	0.044
Share earnings 3. quartile	0.213***	0.032	0.285***	0.016	0.043
Share earnings 4. quartile	0.136***	0.034	0.150***	0.024	0.730
Share income 1. quartile	0.357***	0.047	0.229***	0.015	0.010
Share income 2. quartile	0.268***	0.030	0.302***	0.028	0.417
Share income 3. quartile	0.223***	0.025	0.288***	0.018	0.036
Share income 4. quartile	0.152***	0.033	0.181***	0.025	0.475

Notes: The table reports the complier characteristics (β-estimates of equation (13)) for benefit and wage changes. The change in DI receipt on the right-hand side of equation (13) is instrumented with $RoC_p \cdot A87_t$ for the benefit change and oil employment instrument for the wage change. Earnings and income are measured in $t-3$. The last column reports the p-value for the null hypothesis that marginal entrant characteristics are identical for benefit and wage changes. Levels of significance: *10%, **5%, and ***1%.

Table C.9: Comparison Marginal vs. Average Entrant, Wage Change (Oil Instrument)

	Marginal entrant	Average entrant	P-value
Age	49.43***	45.69***	0.000
	(0.93)	(0.16)	
Family size	2.82***	2.69***	0.110
	(0.09)	(0.02)	
Share female	0.545***	0.585***	0.002
	(0.014)	(0.003)	
Pr(died by age 60)	0.078***	0.092***	0.027
	(0.006)	(0.001)	
Share earnings 1. quartile	0.376***	0.252***	0.000
	(0.020)	(0.003)	
Share earnings 2. quartile	0.189***	0.286***	0.002
	(0.034)	(0.005)	
Share earnings 3. quartile	0.285***	0.262***	0.114
	(0.016)	(0.002)	
Share earnings 4. quartile	0.150***	0.200***	0.028
	(0.024)	(0.005)	
Share income 1. quartile	0.229***	0.212***	0.195
	(0.016)	(0.003)	
Share income 2. quartile	0.302***	0.308***	0.840
	(0.028)	(0.003)	
Share income 3. quartile	0.288***	0.271***	0.335
	(0.018)	(0.002)	
Share income 4. quartile	0.181***	0.210***	0.223
	(0.025)	(0.004)	

Notes: The table reports the complier characteristics (β-estimates of equation (13)) for wage changes and compares the marginal entrant to the average entrant. Earnings and income are measured in $t-3$. The change in DI receipt on the right-hand side of equation (13) is instrumented with the oil employment instrument. The last column reports the p-value for the null hypothesis that marginal entrant characteristics are identical to characteristics of the average entrant. Levels of significance: *10%, **5%, and ***1%.

Table C.10: Comparison of Complier Characteristics, Wage Change (Industry Share Instrument, 1982–2001)

	Benefit change		Wage change		Equality test
	Coeff.	Std. err.	Coeff.	Std. err.	(P-value)
Age	50.57***	0.81	48.76***	2.10	0.422
Family size	3.60***	0.26	3.33***	0.26	0.454
Share female	0.426***	0.038	0.374***	0.051	0.422
Pr(died by age 60)	0.132***	0.017	0.119***	0.018	0.604
Share earnings 1. quartile	0.377***	0.047	0.449***	0.059	0.338
Share earnings 2. quartile	0.274***	0.026	0.283***	0.073	0.913
Share earnings 3. quartile	0.213***	0.032	0.157***	0.035	0.234
Share earnings 4. quartile	0.136***	0.034	0.112	0.092	0.805
Share income 1. quartile	0.357***	0.047	0.431***	0.052	0.294
Share income 2. quartile	0.268***	0.030	0.234***	0.089	0.719
Share income 3. quartile	0.223***	0.025	0.186***	0.031	0.357
Share income 4. quartile	0.152***	0.033	0.149	0.092	0.978

Notes: The table reports the complier characteristics (β-estimates of equation (13)) for benefit and wage changes for the time period 1982-2001. The change in DI receipt on the right-hand side of equation (13) is instrumented with $RoC_p \cdot A87_t$ for the benefit change and the industry shift-share instrument for the wage change. Earnings and income are measured in $t-3$. The last column reports the p-value for the null hypothesis that marginal entrant characteristics are identical for benefit and wage changes. Levels of significance: *10%, **5%, and ***1%.

Table C.11: Comparison of Complier Characteristics, Wage Change (Oil Instrument, 1982–2001)

	Benefit change		Wage change		Equality test
	Coeff.	Std. err.	Coeff.	Std. err.	(P-value)
Age	50.57***	0.81	50.62***	0.83	0.967
Family size	3.60***	0.26	2.86***	0.11	0.008
Share female	0.426***	0.038	0.522***	0.022	0.029
Pr(died by age 60)	0.132***	0.017	0.088***	0.011	0.027
Share earnings 1. quartile	0.377***	0.047	0.416***	0.017	0.437
Share earnings 2. quartile	0.274***	0.026	0.166***	0.039	0.022
Share earnings 3. quartile	0.213***	0.032	0.255***	0.016	0.242
Share earnings 4. quartile	0.136***	0.034	0.163***	0.026	0.533
Share income 1. quartile	0.357***	0.047	0.240***	0.018	0.020
Share income 2. quartile	0.268***	0.030	0.321***	0.036	0.260
Share income 3. quartile	0.223***	0.025	0.240***	0.029	0.661
Share income 4. quartile	0.152***	0.033	0.199***	0.028	0.271

Notes: The table reports the complier characteristics (β-estimates of equation (13)) for benefit and wage changes for the time period 1982-2001. The change in DI receipt on the right-hand side of equation (13) is instrumented with $RoC_p \cdot A87_t$ for the benefit change and the oil employment instrument for the wage change. Earnings and income are measured in $t-3$. The last column reports the p-value for the null hypothesis that marginal entrant characteristics are identical for benefit and wage changes. Levels of significance: *10%, **5%, and ***1%.

D Additional Results for Welfare Analysis

Table D.12: Welfare Analysis, Years 1982-2001

A. Δ DI take-up per $1,000				
	DI benefits ($\partial DI/\partial b$)	Lifetime income ($\partial DI/\partial w$)		
		Industry share	Oil employment	Oil price
Coeff. estimate	0.208***	-0.082***	-0.084***	-0.052***
	(0.024)	(0.024)	(0.023)	(0.010)
B. Welfare impacts				
	Multiplier $\left(1+\frac{\varepsilon_{DI,b}}{1-DI}\right)$	Insurance value $\left(-\frac{\partial DI/\partial b}{\partial DI/\partial w}\right)$		
		Industry share	Oil employment	Oil price
Estimate	1.591***	2.542***	2.485***	4.020***
	(0.069)	(0.810)	(0.733)	(0.881)
P-value: estimate = 1	0.000	0.057	0.043	0.001
P-value: multiplier = ins. value		0.227	0.207	0.004

Notes: The table combines the estimates from column 1 of Table 1 and Panel A of Table C.4 to construct the multiplier and insurance value of higher DI benefits. Levels of significance: *10%, **5%, and ***1%.

Table D.13: Welfare Analysis, Oil and Gas Provinces

A. Δ DI take-up per $1,000				
	DI benefits ($\partial DI/\partial b$)	Lifetime income ($\partial DI/\partial w$)		
		Industry share	Oil employment	Oil price
Coeff. estimate	0.160***	-0.035**	-0.031**	-0.043***
	(0.022)	(0.016)	(0.016)	(0.007)
B. Welfare impacts				
	Multiplier $\left(1+\frac{\varepsilon_{DI,b}}{1-DI}\right)$	Insurance value $\left(-\frac{\partial DI/\partial b}{\partial DI/\partial w}\right)$		
		Industry share	Oil employment	Oil price
Estimate	1.515***	4.531**	5.073*	3.720***
	(0.071)	(2.140)	(2.638)	(0.800)
P-value: estimate = 1	0.000	0.099	0.123	0.001
P-value: multiplier = ins. value		0.155	0.174	0.004

Notes: The table combines the estimates from column 2 of Table B.2 and Panel B of Table C.4 to construct the multiplier and insurance value of higher DI benefits. Levels of significance: *10%, **5%, and ***1%.