Risk Protection and Redistribution in the Design of Social Insurance

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Abstract

We study how heterogeneity in employment risk and exposure to income loss shapes the social value of expanding unemployment insurance (UI) generosity. We estimate individual willingness-to-pay (WTP) for UI reforms, capturing both insurance value and cross-subsidization. WTP to expand U.S. UI peaks among lower-income households, who are subsidized by others under the current system. For a \$1 increase in UI benefit expenditure, social gains are \$0.43 from risk protection and \$0.10 from cross-subsidization, while incentive costs are \$0.71. Of the resulting \$0.54 resource-reallocation gain, 69% reflects efficiency, 20% cross-income and 11% within-income redistribution.

Keywords: social insurance, redistribution, risk protection, consumption, unemployment insurance

JEL classification: E21, E24, H23, H31, H50, I38, J64, J65

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1. Introduction

One central role of government is to insure households against adverse shocks. In the United States, social insurance programs protect households against a wide range of risks, including unemployment, disability, and poor health. Roughly 36 per cent of the federal government's budget—about $2.4 trillion—is devoted to such programs.[1]

Designing effective programs involves a fundamental trade-off between providing insurance and minimizing behavioral distortions. The canonical framework for analyzing this trade-off uses a small set of empirical "sufficient statistics" to characterize welfare-relevant elasticities and consumption responses (e.g., Baily 1978; Chetty 2006; Chetty and Finkelstein 2013). Under certain conditions, the welfare gain from insurance can be inferred from the average consumption gap between "good" and "bad" states of the world, while the associated incentive cost is summarized by a single behavioral elasticity.

While the benchmark approach offers powerful insights, it relies on average responses and abstracts from heterogeneity in both the value of insurance and the cost of behavioral responses. In reality, individuals differ widely in their exposure to shocks, ability to smooth consumption (Blundell et al. 2008), and responsiveness to incentives (Chetty 2008; Attanasio et al. 2018). Moreover, because premiums are rarely fully risk-adjusted, social insurance programs often involve implicit cross-subsidization across individuals. These differences imply that the welfare consequences of reforms to program generosity can vary substantially across individuals.

This paper studies the value of changing the generosity of unemployment insurance (UI) in the United States, with particular emphasis on how heterogeneity in employment risk and exposure to income loss during unemployment shapes the welfare impact of reforms. We provide novel evidence on individual-level willingness-to-pay (WTP) for UI. A key innovation is that we estimate counterfactual consumption during unemployment for workers who are never observed unemployed in our data, enabling us to construct worker-level WTP. We use these estimates to decompose the welfare effects of reform into components reflecting the social value of surplus redistribution—through risk protection and cross-subsidization—and the incentive cost reflected in fiscal externalities.

We use data from the Panel Study of Income Dynamics (PSID) spanning 1999–2019. This nationally representative panel tracks individuals over time and provides detailed information on labor market outcomes, demographics, and household

[1] We sum Social Security retirement, disability, and survivor insurance; unemployment insurance; and Medicare insurance for those age 65 and older. This share has remained relatively stable over time: Feldstein (2005) reports a value of 37% for 2003.

consumption expenditures. To motivate our analysis, we begin by documenting that workers who experience an unemployment spell have lower equivalized household consumption, individual earnings, and household income *prior* to job loss than those who remain continuously employed. Moreover, even conditional on income, workers who go on to experience unemployment exhibit lower *in-work* consumption. These patterns highlight the redistributive role of unemployment insurance—both across and within income groups.

Our worker-level measure of WTP for UI reform has two components. The first is the individual's *personalized insurance value*, which captures their exposure to employment risk—specifically, the value they place on a marginal transfer of resources from states in which they are employed to states in which they are not. The second is their *net fiscal position* with respect to the reform—that is, the extent to which they are, ex ante, net contributors to or beneficiaries of the program. The insurance value varies across individuals depending on the severity of income loss in unemployment and their capacity for self-insurance. The net fiscal position reflects the gap between an individual's risk and the implicit premium embedded in non-individualized taxes and benefits. When implicit contributions deviate from actuarially fair pricing, this gap gives rise to cross-subsidization—i.e., redistribution of expected resources across individuals.

To measure the insurance component of WTP, we build on the "consumption-based" approach to valuing social insurance, developed by Gruber (1997). We extend this approach by providing new evidence on the full distribution of unemployment-induced consumption declines across workers, rather than focusing only on average effects.

To do so, we adopt clustering regression techniques introduced by Lewis et al. (2024) and used to estimate heterogeneous marginal propensities to consume (see also Boehm et al. 2025). These methods allow us to uncover heterogeneity in consumption responses across latent groups without requiring strong assumptions about which observable characteristics drive this variation. Crucially, the framework allows us to recover counterfactual consumption declines for workers we never observe experiencing unemployment. We estimate an average unemployment-induced consumption drop of 11.7%, broadly in line with prior studies (e.g., Gruber 1997; Hendren 2017; Kroft and Notowidigdo 2016). However, we also find substantial heterogeneity reflecting differences in the capacity for self-insurance: for example, the consumption decline at the 95th percentile is roughly twice as large as the median decline.

The cross-subsidization component of WTP is shaped by the interaction between individual employment risk and the structure of the reform. We measure employment risk by estimating how the incidence of unemployment varies with workers demographic and socioeconomic characteristics (see also Anderson and Meyer 2006). We compare a natural expansion of the U.S. UI system—which replaces a fraction of lost earnings up to a cap—with alternative reforms that expand entitlement either by a flat amount or in proportion to lost earnings with no cap.

We find that WTP to expand the generosity of the current U.S. UI system is highly heterogeneous. Median WTP declines with household income, turning negative around the middle of the distribution. This pattern arises even though all workers value actuarially fair insurance positively. The decline reflects the fact that, on average, workers above the median income level are net contributors who cross-subsidize others. Low-income workers benefit from the system because they face higher employment risk—an effect strong enough to outweigh the link between benefit entitlements and earnings. A flat benefit reform further tilts the WTP distribution in favor of lower-income workers, whereas a proportional expansion dampens the income gradient by increasing entitlements more for higher earners.

We also find substantial heterogeneity in WTP even among workers with similar incomes. For example, within the top income decile, the standard deviation of WTP for expanding the current system is more than half as large as the difference between median WTP at the bottom and top of the income distribution. These reform-specific patterns of winners and losers—both across and within income groups—shape the social value of alternative UI reforms.

We next consider the social value of a budget-balanced expansion of UI generosity. Since this requires interpersonal comparisons, our evaluation depends on the specification of social preferences. In our baseline implementation, we adopt welfare weights that value cross-income surplus redistribution at the marginal cost to the government of achieving equivalent redistribution through the income tax system (Hendren 2020), and place greater value on within-income-group surplus redistribution for individuals with lower long-run equivalized household consumption, which we treat as measuring lifetime income and economic well-being (Poterba 1989; Meyer and Sullivan 2003). This formulation reflects the idea that broad redistribution across long-run income groups could, in principle, be implemented through the income tax and transfer system. By contrast, UI also reallocates resources within income groups, targeting households who both face high employment risk and have low in-work consumption or limited self-insurance capacity—dimensions that are difficult to observe and condition on in the tax code. Our framework isolates the part of UI's redistributive value that is specific to the unemployment margin, over and above what could be achieved by adjusting the income tax schedule alone.

We show that the social value of expanding UI can be decomposed into three distinct components. The first is the value of risk protection, which takes the form of a risk- and welfare-weighted average of individual insurance values, scaled by an aggregate welfare weight. The individual weights reflect heterogeneity in the marginal social value of surplus among program beneficiaries, while the aggregate weight captures their average marginal value relative to those who fund the program. We estimate that the risk-protection component of a budget-balanced expansion of the current UI program is $0.43 per dollar of increased benefit expenditure.

The second component captures the social value of cross-subsidization, which arises independently of insurance value. Absent in representative-agent frameworks, this component can generate welfare gains (or losses) even when individuals are fully insured through other means. For the current UI program, we estimate that the cross-subsidization component associated with increased generosity is $0.10 per dollar of increased benefit expenditure. Combining these first two terms, the total social value of resource reallocation is therefore $0.54. By comparing our baseline welfare weights to two benchmarks—one that preserves the cross-income profile implied by the inverse-optimum weights but assigns equal weight to all individuals at a given income level, and one that uses fully money-metric weights that treat all individuals symmetrically—we further decompose this resource-reallocation gain into efficiency improvements from additional insurance (69%), cross-income surplus redistribution (20%) and within-income redistribution (11%).

The third component reflects the fiscal externality associated with behavioral responses to program generosity. In the canonical model, this is typically summarized by a single elasticity. In our setting, by contrast, the fiscal externality depends on how individuals' behavioral elasticities covary with their risk levels and net fiscal positions (i.e., the gap between taxes paid and benefits received across states). We calibrate the behavioral elasticities using evidence from Chetty (2008). We find that the direct fiscal externality from increased UI payments is $0.38 per dollar of increased benefit expenditure, with an additional cost of $0.33 due to lower income tax receipts. Thus, the total incentive cost of $0.71 per dollar exceeds the $0.54 resource-reallocation gains from risk protection and cross-subsidization.

Our measure of the social value of changing UI generosity depends critically on the design of the reform. Under a flat expansion—where UI entitlements increase uniformly across workers—the value of resource reallocation rises to $0.63 per dollar of increased benefit expenditure, as this form of expansion directs a larger share of resources toward households that are less well-off and more exposed to employment risk. In this case, the welfare gains from increased generosity outweigh the associated incentive costs. These differences across reform types underscore that the social value of UI is likely to vary substantially across countries, due to substantial differences in program design (Spinnewijn 2020). The standard sufficient-statistics implementation effectively evaluates a flat reform under social preferences that are indifferent to how surplus is distributed, yielding a much lower estimated social value of resource reallocation of just $0.38. We show that abstracting from surplus redistribution across heterogeneous individuals meaningfully alters the conclusions of UI policy analysis.

We contribute to a literature that uses a sufficient-statistics approach to study the insurance-incentive trade-off in the design of UI (e.g., Schmieder et al. 2012; Kolsrud et al. 2018; Landais and Spinnewijn 2021).[2] Our contribution is to extend this sufficient-statistics framework beyond a representative-agent setting by combining microdata on consumption and employment risk to recover individual-level WTP and to decompose the social value of UI reforms across heterogeneous workers. We also build on recent work that measures the value of social transfers, including disability insurance (Deshpande and Lockwood 2022), public pensions (Kolsrud et al. 2024), and welfare programs (Rafkin et al. 2023).

Our work also relates to a literature that examines how the correlation between risk and ability shapes the joint design of income taxes and social insurance (e.g., Blomqvist and Horn 1984; Rochet 1991; Cremer and Pestieau 1996; Boadway et al. 2006), as well as research on how involuntary unemployment (Kroft et al. 2020) and the structure of non-linear UI benefit schedules (Ferey 2022) influence optimal tax formulas. This work highlights that, in these models, redistribution across long-run income or ability types can, in principle, be implemented through the non-linear income tax and transfer system, while social insurance instruments such as unemployment insurance play a distinct role by providing state-contingent transfers within types and across employment states. We complement this body of work by empirically measuring the extent to which current UI systems redistribute surplus across heterogeneous workers—both across and within income groups.

The rest of this paper is structured as follows. Section 2 outlines our measure of individual willingness-to-pay for increased UI generosity. Section 3 introduces our data and presents preliminary evidence on negative selection among the unemployed. Section 4 describes our empirical approach for estimating willingness-to-pay and presents the main results. Section 5 develops a normative framework for valuing changes in UI generosity and reports our quantitative analysis. Section 6 concludes.

[2]A complementary strand of the literature employs dynamic structural models to analyze UI design (e.g., Acemoglu and Shimer 1999; Lentz and Tranaes 2005; Krusell et al. 2010) and highlights the redistributive implications of UI policy across heterogeneous workers (e.g., Audoly 2024; Haan and Prowse 2024).

2. Individual's Value of a Reform

In this section, we derive an expression for individual-level willingness-to-pay (WTP) for an expansion in the generosity of a social insurance program. We present a simple framework that captures the salient features of unemployment insurance, as well as other programs such as workers' compensation, disability insurance, health insurance, and insurance against natural disasters. The key sufficient-statistics objects we derive extend to a much richer dynamic environment (Chetty 2006; see Appendix A.1).

We build on the canonical sufficient-statistics setup (see Chetty and Finkelstein 2013) by allowing for a heterogeneous population of individuals exposed to risk. Individual-level WTP is a key input for evaluating the social value of reform, which in our framework depends both on how these WTPs are aggregated across individuals and on the incentive costs generated by behavioral responses (see Section 5).

2.1. Theory

Individual's problem. There is a unit continuum of individuals, indexed by i. Each individual faces uncertainty over two states: high (h) and low (l), for instance, corresponding to being employed and unemployed. Individuals may differ in their preferences, incomes, and the costs associated with mitigating risk. We denote income in the high and low state by y_i and z_i, respectively. The government provides a benefit $\mathcal{B}(y_i)$ to those in the low state and levies a tax $\mathcal{T}(y_i)$ on individuals in the high state. We allow the tax and benefit schedules to be arbitrary functions of high-state incomes.[3]

Consumption in the high state is therefore $c_i^h = y_i - \mathcal{T}(y_i)$, while consumption in the low state is $c_i^l = z_i + \mathcal{B}(y_i)$. Let $u_i^s(c)$ denote state-specific utility from consumption c; we assume $u_i^s(\cdot)$ is increasing, concave and twice continuously differentiable. Let $\Delta c_i \equiv c_i^h - c_i^l$ denote the gap in consumption between the high and low states. For most individuals, we expect $\Delta c_i > 0$.

Individuals can control the probability of being in the high state by undertaking actions, which we model through a scalar e and refer to as effort. We normalize the units of effort so that e equals the probability of being in the high state (and therefore $e \in [0, 1]$). Effort is costly, with cost $\psi_i(\cdot)$, which is increasing and convex.

[3] In our empirical implementation, we also account for the role of family composition in the U.S. tax-and-transfer system. We omit this here to simplify exposition.

The individual solves:

$$V_i = \max_{e \in [0,1]} \; e\, u_i^h(c_i^h) + (1-e)\; u_i^l(c_i^l) - \psi_i(e), \tag{1}$$

where V_i is the maximized value of their expected utility. Individual i's optimal effort choice, e_i, satisfies the first-order condition: $u_i^h(c_i^h) - u_i^l(c_i^l) = \psi_i'(e_i)$, i.e., the individual chooses effort such that the marginal benefit of exerting additional effort—equal to the utility difference in the two states—equals the marginal cost of effort.

Heterogeneity. The individual index i captures arbitrary heterogeneity arising from several sources. First, differences in earning ability—for example due to variation in innate ability, human capital, skills, or labor market opportunities—are reflected in high-state incomes y_i. Low-state income z_i additionally reflects an individual's capacity to self-insure, for example through private savings or spousal insurance. As a result, differences in incomes across states embody both the direct financial loss from entering the low state and individuals' self-insurance capacity. Combined with the influence of the tax and benefit system, these factors generate heterogeneity in the consumption gap Δc_i.

Second, heterogeneity in attitudes toward work and employment opportunities (in the unemployment or disability insurance context) or in the cost of abatement investments (in the health or natural-disaster insurance context) is captured by the effort-cost function $\psi_i(\cdot)$.

Third, variation in the marginal utility of consumption (for example, due to household composition) and in attitudes toward risk is captured by individual-specific utilities, $u_i^s(\cdot)$. This formulation also allows for variation in state dependence across individuals, which may reflect differences in the opportunity cost of time or the availability of consumption substitutes.

Each of these factors influences the individual's optimal effort choice and hence generates heterogeneity in risk levels e_i.

Arbitrary reforms. We consider the effects of an arbitrary marginal reform to the benefit and tax schedules.[4] We parameterize the reform by $(db, d\tau)$, such that the implied state-contingent budget constraint in each state shifts according to:

$$dc_i^l = \phi_{\mathcal{B}}(y_i)\, db, \qquad dc_i^h = -\phi_{\mathcal{T}}(y_i)\, d\tau, \tag{2}$$

i.e., the resources available in the low and high states change by $\phi_{\mathcal{B}}(y_i)db$ and $-\phi_{\mathcal{T}}(y_i)d\tau$, respectively, where $\phi_{\mathcal{B}}(y_i)$ and $\phi_{\mathcal{T}}(y_i)$ are positive functions of high-state income y_i. While these functions allow for reforms that vary arbitrarily by income, in practice reforms are typically simple functions of income—for example, a flat benefit increase ($\phi_{\mathcal{B}} = 1$) or proportional benefit increase ($\phi_{\mathcal{B}} = y_i$), funded by a flat tax adjustment ($\phi_{\mathcal{T}} = 1$). The structure of the reform is central to determining individual willingness-to-pay, because it governs both who receives additional resources in each state and who finances them.

Willingness-to-pay. To define the individual-level WTP for a social insurance expansion, we consider the reform

$$d\theta \equiv \left(db = 1,\ d\tau = \frac{\int_i (1 - e_i)\phi_{\mathcal{B}}(y_i) di}{\int_i e_i \phi_{\mathcal{T}}(y_i) di} \right).$$

This reform raises benefits and adjusts the tax schedule just enough to offset the resulting mechanical budgetary cost.

Let $\frac{dV_i}{d\theta}$ denote the impact of this reform on individual i's expected utility. We define their willingness-to-pay for this reform as:

$$\begin{aligned} \text{WTP}_i &\equiv \frac{1}{(1 - e_i)\phi_{\mathcal{B}}(y_i)} \frac{1}{u_i^{h\prime}(c_i^h)} \frac{dV_i}{d\theta} \\ &= \underbrace{\left(\frac{u_i^{l\prime}(c_i^l) - u_i^{h\prime}(c_i^h)}{u_i^{h\prime}(c_i^h)} \right)}_{\text{insurance}_i} + \underbrace{\left(1 - \frac{e_i \phi_{\mathcal{T}}(y_i)}{(1 - e_i)\phi_{\mathcal{B}}(y_i)} \Big/ \frac{\int_{i'} e_{i'} \phi_{\mathcal{T}}(y_{i'}) di'}{\int_{i'} (1 - e_{i'}) \phi_{\mathcal{B}}(y_{i'}) di'} \right)}_{\text{cross-subsidization}_i} . \end{aligned} \tag{3}$$

[4] Focusing on marginal reforms has the advantage that, by the envelope theorem, the first-order impact of the reform on expected utility is fully captured by the change in the state-contingent budget constraint, evaluated at the pre-reform optimum. Throughout, when we write the effect of the reform in terms of dc_i^l and dc_i^h, this should be understood as the induced change in state-contingent consumption possibilities rather than as the policy directly choosing consumption. As a result, willingness-to-pay can be characterized without a full specification of the individual's optimization problem, and WTP in our simple framework aligns with that in a richer dynamic model (see Appendix A.1). The expression can also be interpreted as a first-order approximation to a non-marginal reform, or used to construct higher-order approximations by integrating along the path of a larger reform (e.g., see Kleven 2021).

The scaling in the first line expresses the utility gain from the reform per $1 increase in expected benefit payments, measured in units of the marginal utility of an additional dollar in the high state.

The second line decomposes WTP into two conceptually distinct components. The insurance component captures the value to the individual of transferring a marginal unit of consumption from the high to the low state, valued at their own marginal utilities. This reflects the individual-specific insurance value of social insurance and is likely to be positive for most individuals.

The cross-subsidization component captures how the individual's net expected transfer under the reform compares to that under an actuarially fair adjustment. Under actuarial fairness, the individual would face an expected tax increase of $e_i\,\phi_{\mathcal{T}}(y_i)$ to fund an expected benefit increase of $(1-e_i)\,\phi_{\mathcal{B}}(y_i)$. In practice, however, the actual reform typically departs from actuarial fairness. The cross-subsidization term reflects this deviation: it is positive for individuals whose actuarially fair price is higher than the implicit pooled price embedded in the reform, meaning they are net beneficiaries, and negative for those whose fair price is lower, implying they are net contributors. In other words, it measures whether the individual gains or loses from redistribution relative to a fair-pricing benchmark.

Discussion. Our willingness-to-pay decomposition is related to the expression in Finkelstein et al. (2019), who show that WTP for a marginal expansion of Medicaid can be decomposed into a "pure insurance" term—capturing the covariance between marginal utility and insured medical spending across health states—and a "pure transfer" term—representing the expected reduction in medical spending due to a lower out-of-pocket price. A key difference in our setting is that the pattern of transfers, or cross-subsidization, arises from departures from actuarial fairness embedded in the structure of the tax and benefit reform.

An alternative interpretation of our decomposition is that the cross-subsidization component corresponds to the expected net transfer from the reform when it is valued at a common marginal utility across states, while the insurance component captures additional value that arises because marginal utility is higher in the low state than in the high state. In this sense, the insurance term measures how much the *ex ante*, risk-adjusted value of the reform exceeds its risk-neutral, ex post value. This interpretation aligns with Lieber and Lockwood's (2019) decomposition of the individual-level value of an in-kind transfer.[5]

[5]Both Finkelstein et al. (2019) and Lieber and Lockwood (2019) consider settings with potentially many states of the world and convert utility to monetary terms by scaling by the *average* marginal utility. In contrast, because our empirical application focuses on unemployment insurance, we focus on two states of the world and scale by the marginal utility in the high-income (employment) state.

2.2. Implementation

The individual-level insurance term depends on the gap in marginal utility of consumption between the low and high states. In our baseline implementation, we assume that marginal utility is state independent and that individuals share a common local curvature of utility at their high-state consumption level.

ASSUMPTION 1. *Marginal utility is state independent (i.e., ${u^{s}}'_i(.) = u'_i(.)$ for $s \in \{l, h\}$) and relative risk aversion at high-state consumption is homogeneous across individuals (i.e., $-\frac{u''_i(c^h_i)c^h_i}{u'_i(c^h_i)} = \gamma$).*

Below we discuss the implications of relaxing both state independence and the assumption of common relative risk aversion. Following Gruber (1997), we adopt a consumption-based approach to measuring the marginal-utility gap. Under Assumption 1, a second-order (quadratic) approximation to the utility function around c^h_i yields:

$$\frac{u'_i(c^l_i) - u'_i(c^h_i)}{u'_i(c^h_i)} \approx \gamma \frac{\Delta c_i}{c^h_i}, \tag{4}$$

so that the individual insurance component of WTP can be expressed as the product of their percentage consumption decline between the high and low states and the coefficient of relative risk aversion. Our empirical strategy estimates the unconditional distribution of these consumption declines across the population.

Cross-subsidization depends both on state-specific probabilities and on the structure of the reform under consideration. In our application, we estimate how unemployment probabilities vary across individuals, and consider a range of realistic reforms that differ in how benefit and tax adjustments vary with income.

Discussion. Our implementation adopts a sufficient-statistics approach, using estimates of individual-level insurance value and employment risk to construct willingness-to-pay. We estimate heterogeneous consumption responses to job loss using panel data on earnings and nondurable consumption, extending the tradition of Gruber (1997) as recently employed in the context of UI (e.g., Kolsrud et al. 2018; Ganong and Noel 2019) and other social insurance programs (e.g., Meyer and Mok 2019; Kolsrud et al. 2024). Under Assumption 1, these estimates reveal the insurance component of willingness-to-pay for UI.

We estimate employment risk based on how job loss varies with demographics and a measure of permanent income constructed from a multi-year history of earnings. We interpret the resulting probabilities as summarizing heterogeneity in exposure to employment risk in our sample, rather than as a full structural model of the joint dynamics of wages, assets, and search.

Other approaches to valuing social insurance programs—based on search responses (Chetty 2008; Shimer and Werning 2008), differences in marginal propensities to consume across states, or revealed-preference from supplemental UI purchases (Landais and Spinnewijn 2021)—entail different assumptions and institutional and data requirements. In principle, any such method for recovering the insurance value of UI could be used in place of, or alongside, our consumption-based approach to implement equation (3) and the subsequent normative analysis; in this paper we focus on the consumption-based method because it can be applied consistently in the U.S. setting we study.

3. Data and Setting

Our empirical analysis focuses on unemployment insurance (UI). In the United States, the federal government sets broad guidelines on UI eligibility and benefit generosity, but programs are administered at the state level, with states determining benefit duration, generosity and eligibility criteria (see Von Wachter 2019 for a discussion). Benefit payments and entitlement are computed from workers' quarterly employment and earnings histories. Outside of temporary federal extensions during recessions, the modal duration offered by states is six months. Eligible workers receive payments in proportion to their past earnings up to a maximum threshold.

The U.S. UI system is funded through payroll taxes that are legally incident on employers. The federally mandated minimum for taxable earnings is $7,000 per year. While most states apply a higher threshold, in practice the tax often functions more like a head tax on the number of workers than a proportional earnings tax. A distinctive feature of U.S. UI finance is that tax rates vary across employers due to partial experience rating (see Guo and Johnston 2021 for discussion). Recent evidence suggests that firms are unable to pass the firm-specific component of UI taxes through to wages (Guo 2024). As a result, in our empirical quantification of a UI expansion, we assume that only the common component of UI taxes is incident on wages.[6]

[6] Guo (2024) shows the firm-specific component instead induces employment responses. See Spaziani (2024) for analysis of how experience rating alters the risk borne by employers. Anderson and Meyer (2000) similarly find that firm-specific components are largely incident on firms, though they emphasize that employment responses also affect layoffs and subsequent UI claims (and their denials).

Sample. We use data from the Panel Study of Income Dynamics (PSID), a nationally representative survey of the U.S. population. The PSID has surveyed households since 1968, collecting information on demographics and labor market outcomes including unemployment, earnings, and wages. We use data from the 1999–2019 waves of the biennial "new" PSID, which includes comprehensive information on consumption expenditures.[7]

We drop households for which the reference person does not report educational attainment, and restrict the sample to reference persons aged between 23 and 60. This excludes individuals whose labor supply is likely driven by education and retirement decisions. To proxy for UI eligibility, we omit households where the reference person never reports being employed, as well as periods immediately following a spell out of the labor force. This yields 55,671 observations (15,270 families observed for an average of 3.7 waves) of which 3,331 observations correspond to the reference person being unemployed (2,211 families for an average of 1.5 waves). For families that experience changes in the number of adults (such as divorce, marriage, or death) we construct panel identifiers for demographically stable units and omit the year of the change. Throughout, we focus on the labor market status of the reference person, whom we refer to as the "worker."

Key variables. In our main analysis, we define unemployment based on the employment status reported by the worker at the time of the survey interview. We measure worker earnings using reported wages when employed, and define household income as the sum of worker and spousal wages. For each worker, we also construct a measure of permanent income based on the worker-level fixed effect in a log wage regression, conditional on life-cycle wage growth.

We use a comprehensive measure of expenditures on non-durable consumption and services, excluding spending on health care and housing.[8] To account for differences in household composition, we equivalize consumption expenditures using the square root of household size. We convert all financial variables to 2019 dollars. Summary statistics, details on sample selection, and the construction of key variables, including our measure of worker permanent income, are provided in Appendix B.

[7] Before 1999 the PSID was collected annually, but it included limited data on consumption outside of food expenditures. Both Gruber (1997) and Hendren (2017), for example, use this food expenditure data in studies of unemployment insurance.

[8] There are some differences over time in how consumption components are measured, including expansion of measured items and changes in the aggregation or disaggregation of certain categories. Our analysis uses the broadest set of consumption categories available in each wave. We include survey-wave fixed effects in our econometric specification to account for changes in the coverage of our consumption measure.

Preliminary evidence. In the next section, we document heterogeneity in willingness-to-pay for UI. In Section 5, we undertake a normative analysis, comparing a welfare-weighted aggregation of willingness-to-pay to the program's incentive costs. To motivate this analysis, we begin by presenting descriptive evidence on the extent of negative selection among the unemployed. On average, consumption among the unemployed is about three-quarters that of the employed. This reflects both negative selection—individuals who experience unemployment tend to have lower consumption levels even while working—and the direct consumption declines associated with unemployment. In this section, we focus on measuring the extent of negative selection (presenting estimates of unemployment-induced consumption declines in the next section).

Specifically, we examine how household resources during periods of employment differ between individuals who ever experience unemployment and those who do not. By restricting attention to spells in which the worker is employed, this analysis avoids mechanically induced reductions in earnings and associated consumption responses during unemployment spells.

To operationalize this, we estimate:

$$\ln y_{i,t} = \delta D_i + \beta X_{i,t} + u_{i,t}, \tag{5}$$

where $y_{i,t}$ is a measure of household resources, D_i is an indicator for whether worker i (the reference person) has ever experienced an observed unemployment spell,[9] and $X_{i,t}$ is a vector of controls that includes a cubic in the worker's age, indicators for household composition and year fixed effects.[10] The coefficient of interest δ, captures the difference in household resources associated with ever experiencing unemployment.

Figure 1 reports estimates of δ and associated 95% confidence intervals for four measures of resources: household consumption, the reference person's labor income, household labor income, and worker permanent income. For each outcome, we report three specifications: (i) without controls, (ii) with controls, and (iii) excluding post-unemployment periods to eliminate the effects of potential labor-market scarring.[11] Across all specifications, households whose reference person ever experiences unemployment have lower consumption, individual income, and household income *while in work*, as well as lower permanent income,

[9] Specifically, $D_i = 1$ if worker i reports an unemployment spell at any survey interview, including spells occurring between waves.

[10] The presence of negative selection is robust to additionally controlling for education and race (see Appendix D.1).

[11] For permanent income, differences reflect variation in the composition of workers observed before and after unemployment.

than households whose reference person is never unemployed.

To the extent that social insurance programs reallocate resources across income groups, this redistribution shapes their social value. However, redistribution across income levels can, in principle, be achieved or offset through adjustments to the tax and transfer system. By contrast, reallocating resources within income groups—across households with the same current income but different levels of consumption—is much harder to implement through standard tax and transfer policies. Moreover, because consumption may better reflect permanent income and longer-run economic well-being, redistribution based on consumption differences may be especially valuable. The green marker in Figure 1 shows that, even conditional on a flexible function of households income, households that experience unemployment have lower consumption than those that do not.

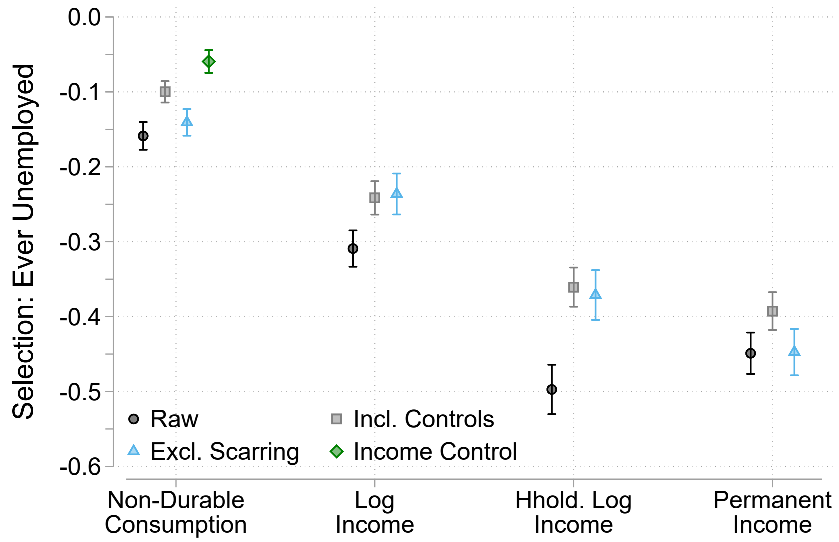


FIGURE 1. Negative Selection Among the Unemployed

Notes: Authors' calculation from PSID data. All regressions drop periods during which the reference person is unemployed. The "excl. scarring" specification further restricts the sample to periods before the first observed unemployment spell. The "raw" specification includes no controls. All other specifications additionally control for a cubic function of age, indicator variables for family size and marital status, and year fixed effects. "Income controls" additionally controls for a cubic function of household income.

4. Measuring Willingness-to-Pay

In this section, we first outline how we measure heterogeneity in unemployment-induced consumption declines and employment risk, which, along with the coefficient of relative risk aversion determine worker risk exposure. In each case, we detail our estimation strategy, present results, and compare them to findings in the existing literature. We then use these estimates to compute individual willingness-to-pay for UI reform.

4.1. Consumption Declines

Method. The standard approach to recovering consumption declines (e.g., Gruber 1997) estimates an average effect of unemployment, conditional on being employed in the preceding period ($t-2$ in our case), using the following specification

$$\begin{aligned} \Delta_{i,t}^{FD} &\equiv \ln\left(c_{i,t}\right) - \ln\left(c_{i,t-2}\right) \\ &= \delta_0 + \delta_1 U_{i,t} + \beta X_{i,t} + \varepsilon_{i,t}, \end{aligned} \tag{6}$$

where $U_{i,t}$ is an indicator variable denoting whether the worker is unemployed in period t[12] and $X_{i,t}$ are demographic controls. This is specified in first differences to allow for arbitrary permanent individual-level heterogeneity in consumption levels. The parameter of interest, δ_1, measures the average proportional consumption decline that results from becoming unemployed. The identification assumption underpinning this strategy is that the trend in consumption for the employed acts as a valid counterfactual for the evolution of in-work consumption of the unemployment had they remained employed.

To estimate the distribution of consumption declines, we specify a finite mixture approximation to the true distribution (e.g., Heckman and Singer 1984; Keane and Wolpin 1997). Our approach follows Lewis et al. (2024), who estimate the distribution of marginal propensities to consume out of tax rebates. We assume each worker belongs to one of a finite number of latent types and augment the specification in equation (6) as follows:

$$\Delta_{i,t}^{FD} = \sum_{k \in K} \mathbb{1}\left[k(i) = k\right]\left(\delta_0^k + \delta_1^k U_{i,t}\right) + \beta X_{i,t} + \varepsilon_{i,t}. \tag{7}$$

We include group-specific intercepts δ_0^k, so that δ_1^k can be interpreted as the consumption decline upon unemployment for group k. We assume that, conditional on group, controls, and unemployment status, the error term is mean zero.

This estimating equation has parallels with the reduced-form in Patterson (2023), who uses unemployment shocks to instrument for income declines. However, an important difference is that we model heterogeneity across unobserved latent types. To implement this, we use a Gaussian mixture linear regression (see Quandt 1972), which assumes that the errors are normally

[12] As in Gruber (1997) and Hendren (2017) we measure unemployment at the time of the interview.

distributed with group-specific variances.

The key identifying assumption in equation (7) is that, conditional on group, the evolution of consumption for the employed acts as a valid counterfactual for the unobserved trend in consumption had the unemployed remained employed. This assumption mirrors that underlying equation (6), but is applied within group. Latent types are identified by grouping households with similar distributions of consumption growth in each state. We estimate group membership jointly with the other parameters (see Appendix C).

We recover estimates of the group-specific parameters, the common effect of covariates, the group-specific variance of the regression residuals, and the unconditional probability of group assignment (π^k for each group k) by maximum likelihood estimation.

To obtain worker-level predictions of consumption declines, we compute mean squared error minimizing "posterior" probabilities of group assignment, $\hat{\pi}_i^k$, at the individual level, which depend on our parameter estimates and the individuals' sequence of observed regressors and outcome variables. We use these to simulate the value of unemployment-induced consumption declines within sample with $\Delta c_i / c_i$ as $\sum_{k \in k} \hat{\pi}_i^k \hat{\delta}_1^k$. These results are conditional on the specified number of groups G which we select using the Bayesian Information Criterion.

As discussed in Lewis et al. (2024), this mixture regression can be viewed as a form of clustering regression which jointly (i) groups households together that have similar latent consumption responses to unemployment and employment and (ii) provides estimates of the consumption decline within these groups. Thus, our approach to estimating counterfactual consumption declines for those without observed unemployment spells can be viewed as a form of matching-based identification strategy. Intuitively, the approach can be thought of as entailing the following steps. First, we group together households observed in both states based on their consumption growth when employed and their consumption decline when unemployed. We then 'match' households who are always employed to those who experience similar consumption growth when employed, conditional on observables. Then, we assign them the matched household's observed consumption decline for their counterparts. In practice, we simultaneously estimate probabilistic group assignment and parameters, with the posterior weight, $\hat{\pi}_i^k$, serving as a (convex) imputation weight.[13]

Estimates. Table 1 reports our estimates from the finite mixture approach alongside estimates from the homogeneous specification (equation (6)). The homogeneous specification yields an estimate of the average consumption decline of 12.9% at the onset of unemployment.

		Mixture Model		
	Homogenous	Group 1	Group 2	Group 3
$\mathbb{1}\left[U_{i,t}=1\right]$	-0.129^{***}	-0.257^{***}	-0.129^{***}	-0.060^{***}
	(0.009)	(0.015)	(0.006)	(0.006)
Share ($\hat{\pi}^g$)		0.117^{***}	0.486^{***}	0.396^{***}
		(0.028)	(0.018)	(0.021)
Controls				
Age	*Yes*	*Yes*	*Yes*	*Yes*
Year	*Yes*	*Yes*	*Yes*	*Yes*
N	37,260	37,260	37,260	37,260

TABLE 1. Heterogeneity in Consumption Declines ($\Delta c/c$)

* $p<0.10$, ** $p<0.5$, *** $p<0.01$. The first column reports the average consumption decline estimated under the assumption of homogeneous consumption drops (equation (6)). The remaining columns report the consumption declines, δ_1^k, and population shares, π^k, in our Gaussian mixture linear regression framework (equation (7)). In the mixture model standard errors account for uncertainty over both group assignment and parameter estimates conditional on group. In addition to the unemployment indicator, we control for a cubic polynomial in age, a series of year dummies, and the log of the change in family size. Results in Table D.1 provide heteroskedastic robust standard errors for the homogeneous estimate.

The implied average consumption decline from our mixture approach, aggregating the group-specific consumption decline δ_1^k, weighted by π^k, is slightly

[13] Note that if two groups experience similar consumption growth when employed, but different consumption declines (e.g., due to varying abilities to self-insure), our imputation assigns a household with similar consumption growth (who is only observed employed) a consumption decline that is a weighted average of these groups. These weights depend not only on the average consumption growth, but also its variability, which corresponds to the pass through of shocks to consumption (Blundell et al. 2008; see Appendix C). In principle, latent consumption risk may vary over time, and our approach could be extended to allow for a regime-switching model.

smaller at 11.7%.[14] This comparison of averages, however, ignores the large degree of heterogeneity in the exposure to unemployment-induced consumption declines that we find. The group-specific consumption declines we estimate range from 25.7% to 6.0%. The probability mass of the group with the smallest decline, 6.0%, is almost 40%. This group is relatively well insured against employment risk. A second group, with probability mass of 49%, has a consumption drop that equals the homogeneous estimate. The remaining probability mass is accounted for by a group with a much larger consumption decline of 25.7%

Figure 2 shows the distribution of posterior predicted consumption declines across worker, further emphasizing the large cross-sectional heterogeneity in risk exposure. Using the standard homogeneous estimator, we have a mass point at the estimate (the dashed vertical line). Instead, we find that 51% of households exhibit consumption drops that are smaller than the standard homogeneous estimate, but there is a long right tail of households with larger drops. This variation reflects broad differences in the ability of households to self-insure employment risk.

Discussion. Our homogeneous estimate is slightly larger than those reported by the existing literature (e.g., Gruber 1997 and Hendren 2017), which find declines in food expenditures of between 7 and 10%. There are three differences between our approach and those in the existing literature that can together account for this.

First, we measure consumption declines by comparing consumption when unemployed with in-work consumption two years prior. Hendren (2017) documents a statistically significant consumption decline of 2.7% in the year preceding unemployment; combining this with existing one-year horizon estimates reconciles the majority of the gap. Second, we use a broader measure of consumption expenditures. Our results are robust to using a wide range of alternative consumption measures (see Appendix D.2). Third, our sample differs from those used in prior work and is drawn from different survey waves. Using the PSID, East and Kuka (2015) document an increasing trend in the average decline in food expenditures following unemployment, which they attribute to factors other than

[14]Part of this discrepancy is the average consumption declines represent subtly different estimands. In the homogeneous specification $\widehat{\delta}_1$ is an estimate of the average consumption decline among the unemployed (or the average treatment on the treated). In contrast, the unconditional estimate in our mixture model estimates the effect in the population (or the average treatment effect) which is possible due to the additional assumptions on the data generating process we impose. When we aggregate estimates of the individual consumption decline by ex-ante unemployment probabilities, we obtain a drop of 12.3%, which is very similar to the homogeneous estimate.

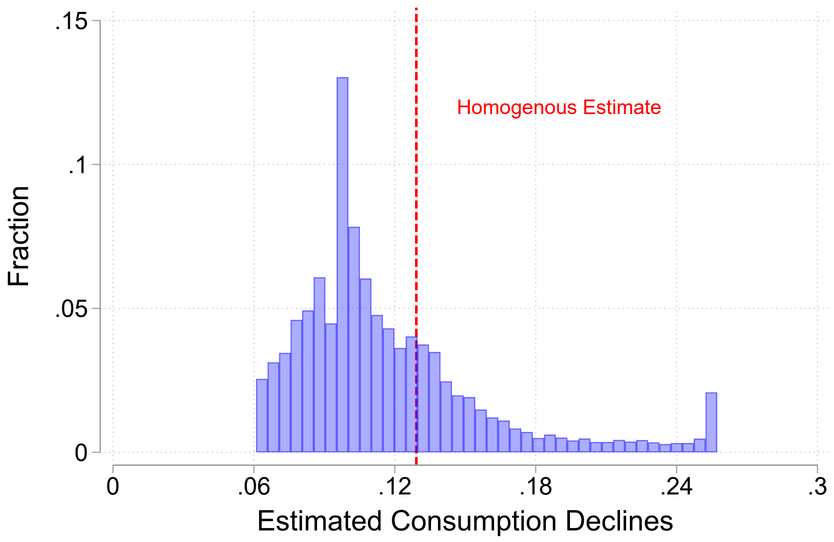


FIGURE 2. Distribution of Estimated Consumption Declines

Notes: Authors' estimates from PSID data. The histogram (light blue bars) plots the worker-level proportional consumption declines constructed using the Gaussian mixture linear regression-estimated parameters and individuals' posterior probabilities for each group, $\hat{\pi}_i^k$. For each household we compute the posterior-weighted effect of unemployment across the discrete group-specific unemployment consumption declines. The sample is defined as in the text. The Bayesian Information Criterion selects K = 3. The homogeneous estimate (red dashed vertical line) is estimated imposing K = 1 in our baseline specification.

changes in the survey design.[15]

Our finding of substantial heterogeneity in exposure to unemployment-induced consumption declines is consistent with a large body of evidence on how households respond to job-loss. Browning and Crossley (2001) find that the marginal propensity to consume out of UI benefits is largest for those without liquid savings. Using direct evidence on binding credit constraints, Crossley and Low (2014) find that 5% of job-losers face binding credit constraints and experience particularly large, welfare-reducing, consumption decline at job loss. They also document smaller, yet substantial, declines among households that are likely to be unconstrained. Ganong and Noel (2019) also document systematic variation in consumption responses by liquid assets, using de-identified banking data. Recent evidence from linked banking and administrative records in Denmark (Andersen et al. 2023) shows that households offset income losses from job-loss primarily by decumulating liquid savings, but

[15]Similar to our analysis, East and Kuka (2015) find only modest differences across consumption measures. They attribute rising declines to a time trend rather than temporal sampling over the business cycle. In addition, our sample includes fewer married white workers compared to earlier work (see Appendix Table B.1 and Table 1 of Kroft and Notowidigdo (2016), who reproduce the estimate from Gruber (1997)). We also explore the sensitivity of our results to alternative definitions of unemployment in the two-year panel and find our results are robust.

also through added worker effects. Finally, both Patterson (2023) and Colarieti et al. (2024) document *marginal* propensities to consume out of unemployment income losses of between 0.5 to 0.6. When workers face different income declines (e.g., married vs. single workers), even constant marginal propensities to consume imply heterogeneous consumption declines. Moreover, Patterson (2023) shows evidence of heterogeneity in these marginal effects by observables.

The key novelty in our analysis is that we construct worker-level estimates of unemployment-induced consumption declines without imposing assumptions about which observables drive this variation. We examine how these estimated consumption declines correlate with proxies for a household's capacity to self-insure and find larger consumption declines among households less likely to benefit from added worker effects (those who do not cohabit) and with low liquid assets (see Appendix D.3).

4.2. Employment Risk

Method. Patterns of program cross-subsidization depend on worker-level risk weights. To estimate these, we use a statistical model of unemployment that recovers rich demographic-specific employment probabilities. This approach infers how risk varies across individuals based on observed realizations of unemployment, sidestepping the need to specify a full model of workers' effort choices.

We specify a logistic regression,[16] where the probability individual i is unemployed at time t depends on the following observed characteristics: a cubic in age, indicators for gender marital status, a full set of education and race indicators, and a cubic in a measure of worker's permanent income.[17]

This yields an estimated employment probability, $\hat{e}_{i,t}$ for each individual and time period. While a non-parametric approach could in principle estimate $e(x) = Pr[U = 0|X]$ more flexibly, the moderate sample size of the PSID motivates our use of a parametric specification.

Estimates. We report average partial effects with 95% confidence intervals in Figure 3 (see Appendix Table D.3 for coefficient estimates). The largest average partial

[16] An alternative approach uses idiosyncratic and subjective probabilities of job loss (Hendren 2017) and job finding (Spinnewijn 2015; Mueller et al. 2021). These measures are not directly available our dataset, and using them would require imputing probabilities based on similar demographic variables.

[17] Our results are nearly identical when we model employment risk directly as a function of estimated consumption declines (see Appendix D.4). This suggests that other factors driving latent heterogeneity in consumption responses do not capture differences in employment risk. Including time effects improves overall fit but does not materially affect the relative risk ranking across demographic groups; in other words, relative employment risks are stable over time.

effects are associated with our measure of permanent income—a one standard deviation increase corresponds to a 4 percentage point decrease in the probability of unemployment.

The results show systematic differences in unemployment rates across demographic groups. These differences are best interpreted as the effect of a given variable for two workers with similar wage histories, as permanent income is included as a control. Younger workers are more likely to be unemployed: a ten-year decrease in age is associated with a 1.4 percentage point increase in the probability of unemployment. We find a modest educational gradient: individuals who did not complete high school are approximately 1 percentage point more likely to be unemployed than those with higher educational attainment. This small effect is largely attributable to the direct role of permanent income. Even after controlling for education and permanent income, we find a substantial racial gap: Black individuals are 1.5 percentage points more likely to be unemployed than white individuals. Conditional on other covariates, we find small effects for gender.

While the simple search model outlined in Section 2 rationalizes these differences through individuals' effort choices, it is important to emphasize that these disparities likely reflect broader factors. These include differences in market thickness, equilibrium congestion externalities, skill-biased labor demand, or outright discrimination—all of which are captured in our model through differences in the cost of exerting effort (i.e. in $\psi_i(.)$).

These patterns are consistent with findings documented in the broader literature. Prior research highlights a positive unemployment rate differential between young and old workers (e.g., Choi et al. 2015); a decline in unemployment rates with increased human capital (e.g., Ashenfelter and Ham 1979; Nickell 1979; Cairó and Cajner 2018); and persistent racial disparities in employment outcomes (e.g., Lang and Lehmann 2012) and callback rates (e.g., Bertrand and Mullainathan 2004). While we find a modest employment gap by gender, it is important to note that our estimates pertain to women who are the reference person in the PSID—a group that is disproportionately single. As prior work shows, participation gaps for unmarried men and women are relatively small, (e.g., Borella et al. 2023).

4.3. Risk Aversion and Reforms

Risk aversion. In our framework, the insurance value of a UI expansion to an individual worker is captured by the product of their consumption decline in unemployment and their coefficient of relative risk aversion (see equation (4)). Based on evidence from a meta-survey of the intertemporal elasticity of substitution (Havránek 2015), we set our baseline coefficient of relative risk aversion to $\gamma = 3$.

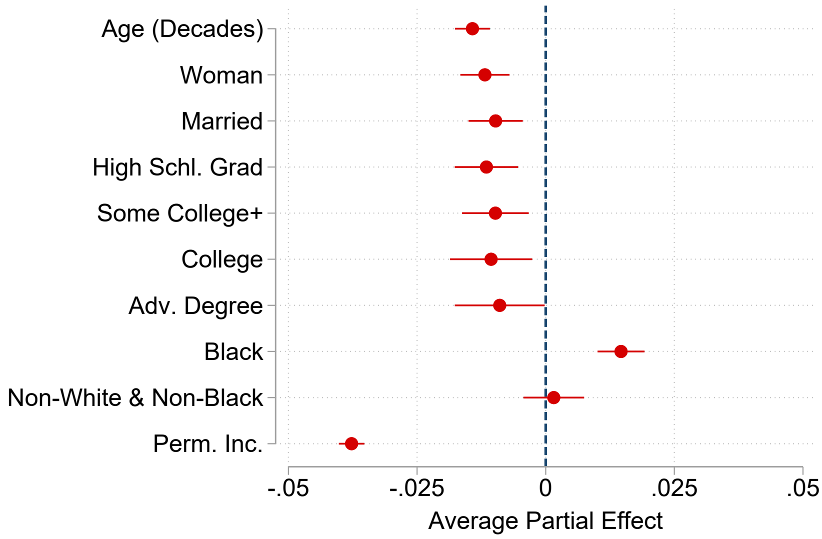


FIGURE 3. Effect of Household Characteristics on Predicted Probability of Unemployment

Notes: Authors' estimates from PSID data. We estimate the logistic regression described in the text using employment status as the dependent variable. We calculate the marginal effects of each covariate on the unemployment probability using parameter estimates and integrate over the empirical distribution of household characteristics. The center of each circle corresponds to the point-estimate of the average partial effect (APE) and the horizontal lines span the 95% confidence intervals. The red dashed line indicates the origin. N=52,996 and standard errors are clustered at the household level. We report APE point-estimates, coefficients estimates and standard errors in Appendix Table D.3.

Reforms. We consider three types of expansion in UI generosity:

(a) Proportional-capped reform: $\phi_{\mathcal{B}}(y) = \min\{y, \kappa\}$, where κ is an earning cap

(b) Flat reform: $\phi_{\mathcal{B}}(y) = 1$

(c) Proportional reform: $\phi_{\mathcal{B}}(y) = y$

All reforms are funded by a flat tax increase, $\phi_{\mathcal{T}}(y) = 1$.

Reform (a) is designed to reflect the most natural expansion of the current U.S. UI system, which offers proportional UI benefits up to an earning cap—approximately at the 41st percentile of the earnings distribution—and is funded by a payroll tax that, in most states, functions as an employment tax. Reform (b) considers a flat benefit expansion for all workers, aligning with the standard reform considered in the UI literature. Reform (c) represents a proportional expansion of UI benefits with no cap, meaning benefits rise in proportion to income for all workers. The interaction between the shape of these reforms and the distribution of employment risk determine the extent of cross-subsidization inherent in each policy.

4.4. Distribution of Willingness-to-Pay

In Figure 4, we summarize the distribution of willingness-to-pay by reform type. Panel (A) shows the insurance component of WTP, which is common across reforms and always positive. Panels (B)-(D) show total WTP for each reform. Each panel plots how the median, interquartile range, and interdecile range of worker-level WTP vary across percentiles of the household income distribution.

Panel (A) shows that WTP for actuarially fair insurance declines with household income and displays substantial variation conditional on income. Among households in the bottom income decile, the median WTP is $0.38 per dollar of benefit expansion—over 20% higher than the median among households in the top decile. This difference is even more pronounced when comparing the upper tail of the WTP distribution within each group: the 90th percentile of WTP in the bottom income decile is $0.68, 55% higher than that in the top income decile.

Panel (B) focuses on a proportional-capped reform, where WTP reflects both the value of insurance and the cross-subsidization embedded in the reform. The wider vertical scale, relative to panel (A), highlights that much of the WTP variation by household income stems from cross-subsidization. Median WTP declines with income—from $0.87 per dollar in the bottom income decile to -$3.66 in the top. Even conditional on income, heterogeneity remains substantial; for example, among workers in the top income decile, the standard deviation in WTP is $2.60.

Although all reforms share a common insurance component, they differ in their patterns of cross-subsidization and therefore in their resulting WTP distributions. Panel (C) shows that a flat reform tilts WTP in favor of workers in lower-income households: median WTP is $1.10 per dollar in the bottom income decile, falling to -$6.40 in the top. Conversely, a proportional reform (panel (D)) tilts gains toward workers in higher-income households.

Under the flat reform, cross-subsidization arises solely from differences in employment risk: workers with below-average risk subsidize those with higher risk. As employment risk tends to be lower among higher-income households, WTP declines steeply with income. Under the proportional reform, cross-subsidization reflects both employment risk and worker earnings. Conditional on employment risk, higher earners gain more from the expansion; yet the overall decline in WTP with household income suggests that declining risk dominates the earnings effect. WTP for a proportional-capped reform resembles the proportional reform at the lower end of the income distribution and the flat reform at the upper end, reflecting the hybrid structure of the policy.

Individual workers' preference orderings over the three reforms depends on both their employment risk and earnings. These reforms create distinct winners and losers and, as we discuss in the next section, differ in their associated fiscal externalities. Ranking their social value therefore requires aggregating willingness-to-pay and accounting for behavioral responses.

Sensitivity to preference specification. The insurance component of WTP depends on our calibrated value of the coefficient of relative risk aversion and the assumption of state-independent utility.

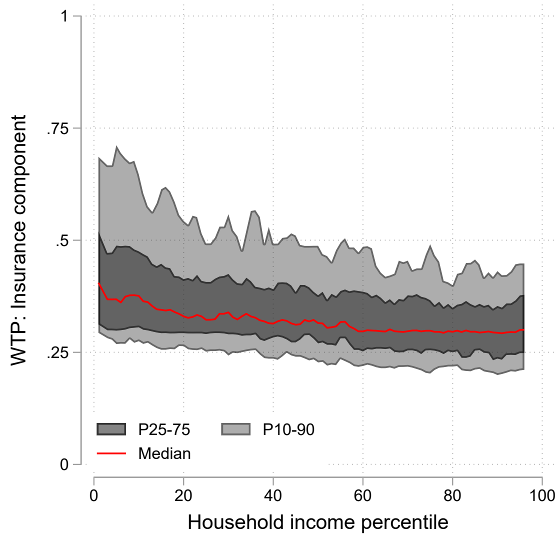
Varying the coefficient of relative risk aversion scales the insurance component proportionally but does not alter the shape of its distribution. We also allow for state-specific consumption prices, capturing the idea that unemployed individuals allocate more time to searching for lower prices. Assuming prices are 1.5% lower for the unemployed—based on evidence from Kaplan and Menzio (2015) and Campos and Reggio (2020)—modestly reduces the insurance component of WTP (see Appendix E.2).

As emphasized by Andrews and Miller (2013), heterogeneity in risk aversion affects the distribution of insurance values only to the extent it covaries with consumption declines. Direct evidence on such heterogeneity is limited.[18] However, as Figure 4 shows, any such effect has a much smaller impact on the distribution of WTP for UI reform than on the insurance component itself, because under realistic reforms the cross-subsidization term plays a dominant role in shaping WTP.

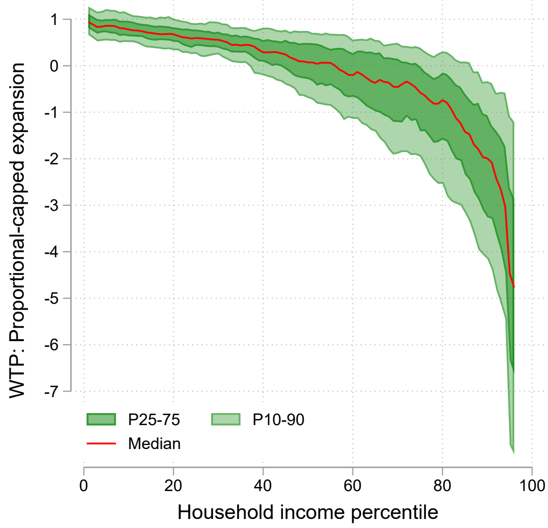
5. Social Value of Reform

In this section, we characterize the social value of expanding a publicly funded insurance program. In the canonical setting, policy trades off the insurance benefit to a representative agent against the cost of distorting incentives, captured by a single elasticity. We show that with heterogeneous individuals, the single representative-agent insurance term is replaced by a welfare-weighted sum of individual willingness-to-pay, which we decompose into the social value of risk protection and cross-subsidization. Moreover, incentive costs now depend on the covariance between individual-level behavioral responses, risk weights, and state-specific net tax liabilities.

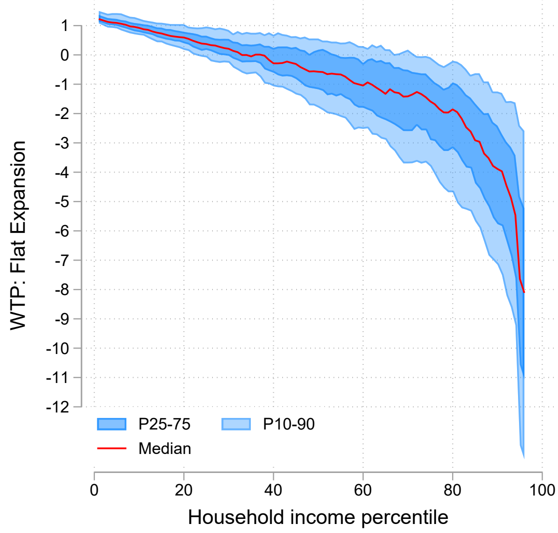
[18] The 1996 PSID survey wave includes survey-elicited measures of individual risk preferences (see Kimball et al. (2009) for more details). These exhibit low covariance with our estimated consumption drops, and therefore imply little change in WTP distributions.



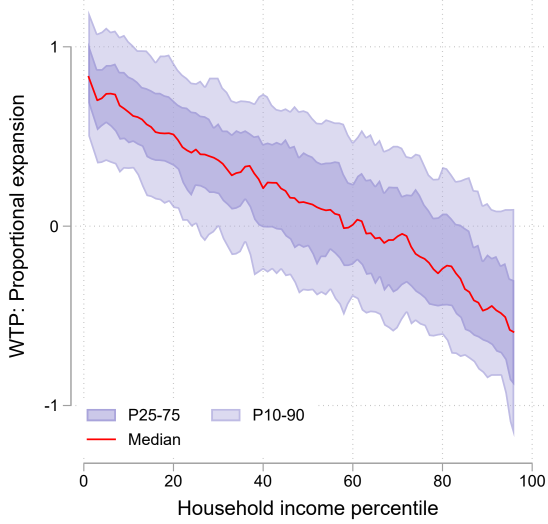
(A) Insurance Component



(B) Proportional-Capped



(C) Flat



(D) Proportional

FIGURE 4. Willingness-to-Pay for UI Expansion by Income and Reform Type
Notes: Authors' estimates from PSID data. Each panel summarizes the distribution of workers willingness-to-pay (WTP) for a UI expansion,reporting the median as well as 10th, 25th, 75th, 90th percentiles. We compute worker-level WTP using equation (3), combining estimates of individual consumption declines and panel-average employment risk, with worker incomes winsorized at \$10,000 and \$200,000. Panel (A) shows results for an actuarially fair reform at the worker level. Panel (B) shows results for a proportional-capped reform reflecting the current U.S. UI structure. Panels (C) and (D) show results for flat and proportional expansions, respectively. Quantiles are smoothed using a uniform rolling window of ± 1 percentiles. Appendix E.1 documents the covariance of individual WTP across different reforms.

5.1. Theory

The aggregate value of a social insurance reform depends both on its fiscal cost and on the weights assigned to the reform's impact on each individual.

To characterize these components, we define risk-weighted expectations in each state. For any variable x_i, we denote its risk-weighted expectation in the low and high state, respectively, as:

$$\mathbb{E}^l[x] \equiv \int_i \left(\frac{(1-e_i)\,\phi_{\mathcal{B}}(y_i)}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})di'} \right) x_i\, di \qquad (8)$$
$$\mathbb{E}^h[x] \equiv \int_i \left(\frac{e_i\,\phi_{\mathcal{T}}(y_i)}{\int_{i'} e_{i'}\,\phi_{\mathcal{T}}(y_{i'})di'} \right) x_i\, di.$$

By construction, the weights in each state sum to one, so $\mathbb{E}^l[x]$ and $\mathbb{E}^h[x]$ are state-specific averages under risk weights proportional to each individual's mechanical contribution to the reform's fiscal impact: in the low state, the expected benefit payment $(1-e_i)\,\phi_{\mathcal{B}}(y_i)$, and in the high state, to the expected tax payment $e_i\,\phi_{\mathcal{T}}(y_i)$.

Budgetary impact. We focus on budget-neutral expansions of social insurance. This implies that the policymaker faces the resource constraint $\int_i e_i \mathcal{T}(y_i)\,di = \int_i (1-e_i)\mathcal{B}(y_i)\,di + \bar{G}$, where $\bar{G}$ denotes an exogenous revenue requirement. The schedules $\mathcal{T}(.)$ and $\mathcal{B}(.)$ summarize the overall tax-and-transfer system in each state: $\mathcal{T}(y_i)$ denotes the net payment to the government in the high state, and $\mathcal{B}(y_i)$ the net transfer received in the low state.

Budget-neutrality implies that the reform $d\theta = (db, d\tau)$, parameterized in equation (2), must satisfy:

$$\frac{d\tau}{db} = \underbrace{\frac{\int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di}{\int_i e_i \phi_{\mathcal{T}}(y_i)\,di}}_{\text{Mechanical effect}} \left(1 + \underbrace{\frac{1}{\int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di} \left(\int_i \frac{d(1-e_i)}{d\theta} (\mathcal{T}(y_i) + \mathcal{B}(y_i))\,di \right)}_{\text{Fiscal externality (FE)}} \right). \qquad (9)$$

The first term captures the tax increase required to cover the mechanical revenue cost of the benefit expansion, holding behavior fixed. The second term reflects the additional tax adjustment needed to maintain budget balance due to behavioral responses—the fiscal externality arising from moral hazard. This arises because higher benefits reduce the marginal return to exerting effort, thereby increasing the likelihood that individuals are in the low state.

Note that while the mechanical effect depends on the design of the reform—

via the functions $\phi_{\mathcal{B}}(.)$ and $\phi_{\mathcal{T}}(.)$—the fiscal externality depends on the pre-reform structure of the tax and benefit schedules $\mathcal{T}(.)$ and $\mathcal{B}(.)$.

We can express the fiscal externality term as:

$$\mathrm{FE} = \mathbb{E}^l \left[\epsilon^{(1-e,b)} \frac{\mathcal{T}(y) + \mathcal{B}(y)}{\mathcal{B}(y)} \right], \tag{10}$$

where $\epsilon_i^{(1-e,b)} \equiv \frac{1}{\phi_{\mathcal{B}}(y_i)} \frac{d(1-e_i)}{db} \frac{\mathcal{B}(y_i)}{1-e_i}$ is the individual-level elasticity of the low-state probability with respect to a benefit increase $\phi_{\mathcal{B}}(y_i)\,db$, financed by the balanced-budget tax adjustment. Equation (10) makes clear that, in general, the fiscal cost arising from moral hazard depends on the covariance between individual effort elasticities and both (i) the risk weights and (ii) state-specific tax and benefit liabilities.

Aggregate value. To capture the aggregate value of the reform, we define a weighted sum of expected utilities:

$$W(\theta) = \int_i \omega_i \, V_i(\theta) \, di, \tag{11}$$

where $\omega_i \geq 0$ is the Pareto weight assigned to individual i. As we discuss below, the inclusion of Pareto weights allows this function to accommodate a broad range of ethical positions regarding the redistributive value of social insurance.

We focus on the aggregate value of a marginal benefit adjustment, parameterized by equations (2) and (9). Analogous to our definition of individual-level willingness-to-pay, we convert the change in social welfare into a money-metric measure by scaling the aggregate utility change $\frac{dW}{d\theta}$ per \$1 balanced-budget increase in the benefit bill, relative to the aggregate utility value of a \$1 decrease in the tax bill. Denoting this money-metric welfare effect by $\frac{dW^{MM}}{d\theta}$, we obtain:

$$\frac{dW^{MM}}{d\theta} = \underbrace{\frac{\mathbb{E}^l[\omega u^{l\prime}(c^l)] - \mathbb{E}^h[\omega u^{h\prime}(c^h)]}{\mathbb{E}^h[\omega u^{h\prime}(c^h)]}}_{\text{Resource reallocation (RR)}} - \mathrm{FE} \tag{12}$$

The first term captures the gain from expanding social insurance that arises from shifting resources from high to low states of the world. It is measured by the percentage gap between the risk- and Pareto-weighted expected marginal utility of consumption in the low and high states. The overall aggregate value of the reform equals this resource-reallocation gain net of the fiscal externality.

Insurance–cross-subsidization decomposition. We define social marginal welfare weights as $g_i \equiv \omega_i u_i^{h\prime}(c_i^h)$. The following proposition decomposes the resource-reallocation term in equation (12) into the societal value placed on insurance—the within-individual, actuarially fair tax transfer of resources between states—and cross-subsidization—the across-individual net expected transfer embedded in the reform.

PROPOSITION 1. *The aggregate value of resource reallocation from a marginal, budget-neutral rise in benefits satisfies:*

$$\mathrm{RR} = \underbrace{\bar{\lambda}}_{\substack{\text{aggregate}\\ \text{welfare weight}}} \mathbb{E}^l \left[\underbrace{\lambda}_{\substack{\text{individual}\\ \text{welfare weight}}} \left(\underbrace{\frac{u^{l\prime}(c^l) - u^{h\prime}(c^h)}{u^{h\prime}(c^h)}}_{\text{insurance}} + \underbrace{1 - \frac{e\,\phi_{\mathcal{T}}(y)}{(1-e)\,\phi_{\mathcal{B}}(y)} \Big/ \frac{\int_{i'} e_{i'}\,\phi_{\mathcal{T}}(y_{i'})\,di'}{\int_{i'} (1-e_{i'})\,\phi_{\mathcal{B}}(y_{i'})\,di'}}_{\text{cross-subsidization}} \right) \right] \tag{13}$$

where $\bar{\lambda} \equiv \frac{\mathbb{E}^l[g]}{\mathbb{E}^h[g]}$ *is the ratio of risk-weighted social marginal welfare weights across the two states, and* $\lambda_i \equiv \frac{g_i}{\mathbb{E}^l[g]}$ *is the welfare weight of individual i, normalized such that the low-state risk-weighted average is 1, i.e.,* $\mathbb{E}^l[\lambda] = 1$*. This can be equivalently expressed as:*

$$\mathrm{RR} = \underbrace{\bar{\lambda}\mathbb{E}^l \left[\lambda \left(\frac{u^{l\prime}(c^l) - u^{h\prime}(c^h)}{u^{h\prime}(c^h)} \right) \right]}_{\substack{\text{risk protection}\\ \text{(within-individual)}}} + \underbrace{(\bar{\lambda} - 1)}_{\substack{\text{cross-subsidization}\\ \text{(across-individuals)}}} \tag{14}$$

Proof: See Appendix F.1.

Equation (13) expresses the welfare gain from social-insurance expansion via resource reallocation as a weighted average of individuals' willingness-to-pay (or surplus), modulated by individual and aggregate social marginal welfare weights. The individual weight λ_i is normalized such that its low-state risk-weighted average is 1 (i.e., $\mathbb{E}^l[\lambda] = 1$). It reflects the social value of increasing individual *i*'s surplus by \$1, relative to the average value among program beneficiaries. The aggregate weight $\bar{\lambda}$ captures the relative weight placed on individuals in the low state compared to those in the high state. It represents the average social value of providing the set of program beneficiaries with an additional \$1 of surplus, relative to the average value of doing so for program funders. This connects closely to the advantageous or adverse selection of beneficiaries relative to funders that we document above. Importantly, the aggregate weight is reform-specific, through the expectation weightings, capturing how the benefit and tax changes vary with income.

Equation (14) expresses the welfare gain from the reform in terms of two components: (1) the social value of insurance, which we refer to as risk protection, and (2) the social value of the cross-subsidization embedded in the implicit pricing of social insurance. This redistribution of net resources across individuals is captured by a single sufficient statistic—the reform-specific aggregate welfare weight $\bar{\lambda}$.[19]

These equations highlight two distinct channels through which social-insurance reform affects aggregate welfare. First, a reform alters the degree of risk protection provided to individuals. All else equal, the value of this protection is higher when there is a stronger positive correlation between individual insurance values and welfare weights. Moreover, its value increases with the relative weight society places on program beneficiaries compared to funders, as captured by the aggregate welfare weight $\bar{\lambda}$. Second, even in the absence of individual-level demand for insurance, the reform reallocates net resources across individuals through cross-subsidization. If individuals in the low state have higher welfare weights than those in the high state (i.e., $\bar{\lambda} > 1$), expanding the program generates positive value purely from this reallocation. This cross-subsidization arises because tax and benefit adjustments typically deviate from actuarially fair terms.

Importantly, because the risk weighting in expectations depends on the design of the reform (via $\phi_{\mathcal{B}}$ and $\phi_{\mathcal{T}}$), both the aggregate insurance and cross-subsidization components are reform specific.

5.2. Implementation

We maintain Assumption 1, which implies state-independent marginal utilities with common local curvature. In addition to individual-level insurance values and risk weights, we require two additional ingredients.

Welfare weights. The social marginal welfare weights g_i encode social preferences over interpersonal utility comparisons. As highlighted by Saez and Stantcheva (2016), these weights accommodate a wide range of normative views about how to compare welfare across individuals. In our baseline implementation, we use the following

[19] Like our willingness-to-pay expression (equation (3)), the decompositions in equations (13) and (14) focus on marginal reforms. This allows us to apply the envelope theorem and ensures consistency with the analogous conditions in a rich dynamic model (see Appendix F.2). The decomposition is general in that it applies both to marginal reforms around the current system—our focus in this paper—and to reforms evaluated at the optimum. As such, it can be combined with a structural model to help characterize optimal policy.

specification:

$$g_i = g_{y_i} \times \tilde{g}_i, \qquad \tilde{g}_i \equiv \frac{(c_i^h)^{-\gamma}}{\mathbb{E}[(c^h)^{-\gamma}|\, y_i]}, \tag{15}$$

where g_{y_i} is the inverse-optimum weight computed in Hendren (2020), and c_i^h denotes long-run in-work (equivalized) consumption. By construction, $\mathbb{E}[\tilde{g} \mid y_i] = 1$, so the inverse-optimum weights g_{y_i} determine the *average* welfare weight placed on each income group, while $\tilde{g}_i$ weights *within* an income group toward individuals with lower long-run consumption.[20]

Under this specification, the social value of transferring \$1 of surplus between individuals at different income levels is governed by the inverse-optimum weights g_y. These weights rationalize indifference to small modifications of the prevailing income tax and transfer system and can be interpreted either as the revealed social preference for redistribution across income groups, or alternatively, as the marginal cost to the government of providing \$1 of surplus to individuals of income y via an adjustment to the tax and transfer system. Using these weights to value cross-income redistribution reflects the fact that equivalent transfers could, in principle, be implemented through the tax and transfer system. As, in practice, most married households file income taxes jointly, we map inverse-optimum weights to household income.

Social insurance reform may also reallocate surplus among individuals with the same level of income. Our welfare-weight specification values this within-income-group surplus redistribution according to $\tilde{g}_i \propto (c_i^h)^{-\gamma}$, implying that, conditional on current income, the social value of providing an additional dollar of surplus to an individual declines with their consumption level. This is motivated by empirical evidence that consumption is a reliable proxy for both lifetime income and economic well-being (e.g., Poterba 1989; Meyer and Sullivan 2003). We measure c_i^h using the panel-average in-work consumption, equivalized by household size to account for differences in needs due to household composition, and winsorize the measure at the 1st and 99th percentiles. While the income tax and transfer system redistributes across current income levels, social insurance reform may deliver additional redistributive value by directing resources to individuals with low consumption levels *within* income groups.

[20] Equivalently, we can write $g_i = \omega_i u_i'(c_i^h)$ in terms of Pareto weights ω_i and marginal utilities $u_i'(\cdot)$. Approximating $u_i'(c_i^h) \approx (c_i^h)^{-\gamma}$ and choosing $\omega_i = \frac{g_{y_i}}{\mathbb{E}[(c^h)^{-\gamma}|\, y_i]}$ implies the social marginal welfare weights in equation (15), $g_i = g_{y_i}\,(c_i^h)^{-\gamma}/\mathbb{E}[(c^h)^{-\gamma} \mid y_i]$.

Fiscal externality. The fiscal externality associated with social insurance reform (equation (10)) depends on the behavioral response elasticity, $\epsilon_i^{(1-e,b)}$, and state-contingent net taxes and benefits, $\mathcal{T}(y_i)$ and $\mathcal{B}(y_i)$.

We calibrate the elasticity using evidence from Chetty (2008), who reports unemployment-duration elasticities separately by quartiles of the wealth distribution.[21] The average elasticity reported in Chetty (2008) aligns closely with the median value reported in the broader literature (see Schmieder and Von Wachter 2016).

We compute state-contingent net tax and benefit schedules using NBER TAXSIM, assuming married households file jointly (see Appendix G.1 for details). This means our measure of the incentive costs of reform accounts for the fiscal impact of reduced worker search effort through both higher UI payments and the resulting decline in income tax revenues.

Discussion. Our characterization of the welfare gains from resource reallocation parallels recent work on the social value of transfer programs. For example, Kolsrud et al. (2024) examine the effects of shifting resources between an early and a late retiree, and show that differences in their consumption levels and welfare weights capture the combined insurance and redistributive effects of reforms. A key difference in our setting is that we model individual-level heterogeneity across the full population of workers, so the social value of reform depends on the joint distribution of insurance values and welfare weights—together with employment risk, which plays a more limited role in the retirement context. As we show below, this is empirically important in our setting. In this respect, our approach also connects to recent work that uses consumption data from the PSID to evaluate the targeting value of disability insurance (Deshpande and Lockwood 2022) and welfare programs (Rafkin et al. 2023).

5.3. Results

Baseline

Row (a) of Table 2 presents our baseline estimate of the social value of increasing UI generosity. These estimates are based on a budget-balanced expansion of the current U.S. UI program, which raises benefits proportionally to the minimum of

[21] The unemployment-duration elasticity is the dynamic analogue of our static elasticity $\epsilon_i^{(1-e,b)}$ (see Appendix F.2). The estimates in Chetty (2008) range from −0.642 in the lowest wealth quartile to 0.016 in the highest (Table 2, pp 204). In our data, we assign each household the corresponding elasticity based on its wealth quartile. We adjust for the exhaustion of UI benefits after 6 months (see Appendix G.1 for details and Schmieder and Von Wachter (2016) for discussion).

a worker's earnings or an earnings cap, funded by a flat increase in payroll taxes. Columns (1) and (2) report the social value derived from risk protection and cross-subsidization, respectively, while column (3) shows the total resource-reallocation effect—the sum of these two components. We estimate that the risk-protection value of a $1 budget-balanced expansion of the UI program is $0.43 and the social value of cross-subsidization is $0.10, yielding a total resource reallocation benefit of $0.54 per dollar of increased benefit expenditure.

These gains must be weighed against the associated fiscal externalities. Column (4) reports the direct fiscal externality from higher UI payments, column (5) reports the additional fiscal externality due to reduced income tax receipts, and column (6) presents the total fiscal cost. We find that accounting for the broader tax-and-transfer system increases the fiscal externality from $0.38 (ignoring changes in income-tax revenues) to $0.71 (including them). Overall, the incentive costs associated with the expansion ($0.71) outweigh the resource reallocation gains ($0.54), implying a net welfare effect of about –$0.17 per dollar of increased benefit expenditure under our baseline welfare weights.

Reform structure

Rows (b) and (c) examine two alternative reforms: a flat benefit expansion (independent of earnings) and an uncapped proportional expansion. Compared to the baseline, the flat expansion is more progressive, allocating a larger share of additional resources to individuals with higher welfare weights. This raises the social value of both risk protection and cross-subsidization; the total resource-reallocation gain from the flat reform is $0.63 per dollar of increased benefit expenditure.

By contrast, the proportional reform is more regressive than the baseline, directing a larger share of resources to workers with earnings above the cap. This reduces the social value of risk protection and cross-subsidization, yielding a lower resource-reallocation value of $0.40 for this reform.

The overall social value of the flat expansion is further enhanced by a reduction in the associated fiscal externality compared to the baseline. This reduction reflects two offsetting forces. On the one hand, the flat reform directs a larger share of the benefit expansion toward lower earners, who tend to exhibit somewhat stronger behavioral responses; this modestly increases the direct fiscal externality. On the other hand, lower earners face smaller income-tax liabilities, which reduces the indirect fiscal externality linked to lost tax revenue—and this effect dominates. For the proportional reform, the pattern is reversed: shifting the incentives to reduce effort toward higher earners who are less responsive, lowers the direct fiscal externality relative to the baseline, but this is more than offset by an increase in the indirect fiscal externality due to their larger tax contributions.

Taken together, these differences across reform structures underscore the importance of incidence in determining the social value of a given policy change.

Social preferences

Row (a) reports our baseline results under the welfare weights in equation (15); rows (d)–(f) illustrate how the social value of the reform varies depending on the weighting placed on individual willingness-to-pay.

Money-metric weights. Under money-metric marginal social welfare ($g_i = 1$), social preferences are indifferent to the distribution of surplus across individuals. One rationale for this approach is given by the Kaldor-Hicks Compensation Principle: if it is theoretically possible to combine the reform with individual-specific lump-sum taxes and transfers to achieve a Pareto improvement, the reform is considered desirable. Since this criterion eliminates any social value attached to redistributive gains, the resource-reallocation value falls to \$0.37, attributable entirely to risk protection (row (d)). Under money-metric weights, the cross-subsidization component of resource reallocation is zero by construction, because a \$1 gain in surplus is valued equally regardless of who receives it. A key drawback of this criterion is that it relies on a lump-sum redistribution scheme that is not practically implementable.

Money-metric weights also clarify the connection between the budget-balanced reform that we focus on and the "unpackaged" reforms consisting of (i) an unfunded benefit increase and (ii) a tax change. UI recipients' average willingness-to-pay for an *unfunded* UI expansion, per mechanical \$1 increase in benefits, is \$1.37—the direct \$1 plus \$0.37 in money-metric risk-protection value implied by the resource-reallocation term in Table 2, row (d).

If we assume that the fiscal externality generated by a balanced-budget UI expansion is entirely due to the benefit expansion rather than the tax rise—so that the fiscal externality from an unfunded benefit expansion equals the fiscal externality in the budget-balanced reform—then each \$1 of higher benefits costs the government \$1.71 once behavioral responses are taken into account (the \$1 mechanical cost plus \$0.71 from behavioral responses). The implied marginal value of public funds (MVPF) for an unfunded UI benefit expansion is therefore \$1.37/\$1.71 $\approx$ 0.8. Under the same assumption, an unfunded payroll tax reduction has a mechanical cost of \$1, generates no fiscal externality, and raises recipients' surplus by exactly \$1, so its MVPF is 1. With money-metric welfare weights, a simple comparison of these MVPFs implies that an additional dollar of government revenue is more valuable when used to cut payroll taxes than to expand UI benefits. When social welfare weights value surplus redistribution, the joint reform can be unpacked in an analogous way into appropriately welfare-weighted MVPFs for UI benefits and for the financing tax (Appendix F.3).

Inverse-optimum weights. Row (e) reports results using inverse-optimum welfare weights, while assuming that social preferences are indifferent to the allocation of surplus *within* income groups—that is, setting the within-income welfare weights $\tilde{g}_i = 1$. Hendren (2020) motivates inverse-optimum weights as an incentive-compatible extension of the Kaldor-Hicks Compensation Principle, which accounts for the distortionary costs of achieving cross-income redistribution through adjustments to the income tax system. Under these weights, the social value of resource reallocation is \$0.48.

Between- versus within-income redistribution. Comparing money-metric weights, inverse-optimum weights, and the baseline results reveals the social value of surplus redistribution.

Comparing row (d) with row (e) highlights the value of cross-income-group surplus redistribution achieved by UI expansion. Moving from money-metric to inverse-optimum weights increases the resource-reallocation benefit of the reform by \$0.11, with over two-thirds of this gain attributable to the direct redistributive value arising from cross-subsidization and the remainder from changes in the welfare-weighted insurance term.

Comparing row (e) with the baseline results in row (a) reveals the additional social value of UI expansion arising from within-income-group surplus reallocation—specifically, between individuals with different levels of long-run in-work consumption. This within-income-group effect further increases the social value of UI expansion arising from resource reallocation by \$0.06, split evenly between risk protection and cross-subsidization. The fact that UI provides insurance against consumption risk—beyond the immediate shock of becoming unemployed—echoes findings from studies of disability insurance (Deshpande and Lockwood 2022) and self-selection into welfare programs among the eligible population (Rafkin et al. 2023).

Of the overall value of resource reallocation of \$0.54 in row (a), \$0.37 (69%) stems from efficiency gains from risk protection, \$0.11 (20%) from cross-income surplus redistribution, and the remaining \$0.06 (11%) from within-income surplus redistribution.

Utilitarian weights. Our use of inverse-optimum weights to value cross-income-group redistribution can be motivated either by the extended Kaldor-Hicks Compensation Principle or by interpreting these weights as reflecting society's revealed preferences for redistribution. One potential objection to the first rationale is that, in practice, social insurance reforms are rarely accompanied by the complex income-tax adjustments required to implement the compensation scheme. The second rationale may be challenged on normative grounds, as these weights may encode preferences for redistribution that are either too strong or too weak.

In row (f), we therefore present results based on utilitarian social welfare weights, defined as $g_i = (c_i^h)^{-\gamma}$. These weights encode stronger preferences for cross-income redistribution than inverse-optimum weights, leading to an increase in both the value of risk protection and the redistributive benefit.

While the precise social value of surplus reallocation from UI reform ultimately depends on the policymaker's chosen welfare-weighting scheme, this analysis demonstrates that its value will generally exceed that implied by a simple unweighted aggregation of money-metric surplus. Specifically, as long as the marginal surplus of (i) lower-income households is valued more than that of higher-income households, (ii) lower-consumption households is valued more than that of higher-consumption households, or (iii) some combination of both, the social value of the reform will exceed that implied by money-metric weights.

Alternative implementations

Standard implementation. The penultimate row of Table 2 reports results for a flat reform evaluated using money-metric welfare weights, approximating the standard Baily-Chetty implementation. The corresponding resource-reallocation value of risk protection is estimated at $0.38 per dollar of increased benefit expenditure. The difference relative to row (b)—which uses baseline welfare weights—reflects the distinction between valuing risk protection as a simple employment-risk-weighted average of individual insurance benefits and applying welfare weights that incorporate the social value of surplus redistribution.

Overall, the comparison between our approach and the standard implementation highlights the importance of accounting for individual-level heterogeneity—and its interaction with the reform design—when evaluating both welfare gains to workers and the incentive costs borne by the government budget.

Average implementation. The final row of Table 2 presents results from an alternative implementation of a flat reform that adapts the approach of Kolsrud et al. (2024), originally applied to retirement benefit reform. This method evaluates equation (14) using *average* consumption among the employed and unemployed populations. It captures redistribution between these groups but abstracts from surplus dispersion within them. In this "average implementation," we likewise value the fiscal externality using the average elasticity, tax liabilities and benefit entitlements of the unemployed. We provide further details in Appendix G.2.

This average implementation yields a resource-reallocation gain of \$0.54 per dollar of increased benefit expenditure—much closer to our estimate for the flat reform (\$0.63) than to the standard implementation (\$0.38). However, by abstracting from the positive covariance between employment risk and welfare weights within the unemployed population, it somewhat understates the social value of resource reallocation. It also overstates the indirect component of the fiscal externality by ignoring the correlation between employment probabilities, behavioral responses, and individuals' state-specific tax positions. The combined effect is a net social value of -\$0.38 under the average implementation, compared with approximately zero under the full implementation.

Overall, the average implementation provides a substantially more accurate picture of the gains from resource reallocation than the standard implementation. However, in our context, it introduces some error, and, importantly for our purposes, is unable to compare the effects of different reforms that generate heterogeneous incidence across the unemployed population.

Cross-sectional implementation. The results in Table 2 exploit the panel dimension of the PSID. Panel data offer two main advantages over cross-sectional data. First, they enable us to estimate worker-level insurance values and WTP, which facilitates a decomposition of the social gains from expanding the generosity of UI into within-individual risk protection and across-individual cross-subsidization. Second, longitudinal information allows us to measure consumption and earnings using panel averages, helping to limit the influence of measurement error. However, if a researcher only has access to cross-sectional data, it is still possible to compute the social value of resource reallocation by directly implementing the social marginal utility gap in equation (12).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Resource reallocation			Fiscal externality			Total effect
	Risk protection	Cross-subsidization	Total (1)+(2)	Direct	Indirect	Total (4)+(5)	(3)-(6)
(a) Baseline	0.43	0.10	0.54	0.38	0.33	0.71	-0.17
	[0.35, 0.53]	[0.09, 0.12]	[0.45, 0.64]	[0.38, 0.39]	[0.31, 0.34]	[0.70, 0.72]	[-0.26, -0.07]
Reform structure							
(b) Flat	0.47	0.16	0.63	0.40	0.19	0.59	0.04
	[0.37, 0.58]	[0.14, 0.18]	[0.53, 0.73]	[0.40, 0.40]	[0.17, 0.21]	[0.57, 0.61]	[-0.06, 0.15]
(c) Proportional	0.40	0.03	0.43	0.36	0.37	0.73	-0.30
	[0.32, 0.49]	[0.01, 0.06]	[0.34, 0.52]	[0.35, 0.37]	[0.35, 0.38]	[0.71, 0.74]	[-0.38, -0.21]
Social preferences							
(d) Money-metric	0.37	0	0.37	=(a)	=(a)	=(a)	-0.34
			[0.30, 0.44]				[-0.41, -0.27]
(e) Inverse-optimum	0.40	0.08	0.48	=(a)	=(a)	=(a)	-0.23
	[0.33, 0.48]	[0.07, 0.08]	[0.40, 0.55]				[-0.31, -0.15]
(f) Utilitarian	0.56	0.38	0.94	=(a)	=(a)	=(a)	0.23
	[0.44, 0.70]	[0.34, 0.42]	[0.82, 1.08]				[0.10, 0.38]
Alternative implementations							
(g) Standard implementation	0.38	-	0.38	0.40	-	0.40	-0.02
(flat reform+money-metric)			[0.31, 0.46]			[0.40, 0.40]	[-0.09, 0.06]
(h) Average implementation	0.43	0.12	0.54	0.40	0.40	0.80	-0.38
	[0.35, 0.52]	[0.11, 0.12]	[0.46, 0.63]	[0.40, 0.40]	[0.39, 0.42]	[0.79, 0.82]	[-0.46, -0.29]

TABLE 2. Risk Protection, Cross-subsidization and Fiscal Externality from UI Expansion

Notes: Authors' estimates based on PSID data. Column (7) reports the total social welfare impact of a UI expansion, decomposed in columns (1)–(6). Row (a) corresponds to an expansion of the current U.S. UI system under baseline social preferences; other rows report results for alternative reform structures or social preferences. All entries are expressed in dollars per $1 increase in benefit expenditure. 95% confidence intervals, based on household-clustered bootstrap samples, are reported in square brackets.

6. Conclusion

In this paper, we study the positive and normative implications of heterogeneity in employment risk and exposure to income loss for the design of unemployment insurance. We find substantial variation in workers' willingness-to-pay for UI, driven by two key factors: (i) differences in the value workers place on insurance, captured by our worker-level estimates of unemployment-induced consumption declines, and (ii) program cross-subsidization, arising from heterogeneity in employment risk and the incidence of reforms. The correlation between this heterogeneity and household well-being—reflected in social welfare weights—shapes the social value of expanding UI. We show that workers who experience unemployment shocks are negatively selected, both in terms of household income and, conditional on income, in terms of consumption while employed. As a result, UI expansion generates surplus redistribution tilted toward less well-off households, through both risk protection and cross-subsidization.

These social gains must be weighed against the distortionary costs of reform. We show that, after accounting for losses from reduced income-tax payments, the net social value of a natural proportional-capped expansion of the current U.S. UI system is negative. In contrast, the net social value of a flat expansion—which directs a larger share of resources toward lower-income households—is approximately zero. Our analysis omits other fiscal externalities, such as the impact of unemployment on disability insurance take-up or future tax payments following re-employment. Quantifying these channels remains an important direction for future work.

Our approach relies on marginal analysis, which has the advantage of sidestepping the need to estimate a full structural model. While it provides insight into the merits of local reforms, it is not designed to characterize optimal policy. A promising avenue for future research is to combine our decomposition with a structural model to assess optimal UI reform.

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Appendix: For online publication

Risk Protection and Redistribution in the Design of Social Insurance

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A. Willingness-to-Pay: Theory

A.1. General model

Here we extend the static model from Section 2 to a richer dynamic setting, following the framework in Chetty (2006) and derive an expression for willingness-to-pay for social insurance reform. As in equation (3), willingness-to-pay decomposes into components reflecting insurance and cross-subsidization. The difference is that, in the dynamic model, these components are evaluated as expectations over the individual's lifetime.

We consider a setting in which time is continuous and individuals live over $t \in [0, 1]$. Let $\varphi_{i,t}$ denote a state variable containing all relevant information up to time t in determining the individual's time t state status (i.e., whether they are in the high or low state) and behavior. $\varphi_{i,t}$ has unconditional distribution $F_{i,t}(\varphi_{i,t})$ at $t = 0$. Assume $F_{i,t}$ is smooth with maximal support φ for all (i, t).

Let $c_{i,t}(\varphi_{i,t})$ denote individual i's time t state-contingent consumption. Let $x_{i,t}(\varphi_{i,t})$ denote M other choices the individual makes (for instance, different dimensions of effort, actions to self-insure like borrowing from family, spousal labor supply decisions and so on). Denote the individual's flow utility function $u_i^s(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}))$ for $s \in \{l, h\}$. Let $\xi_{i,t}(\varphi_{i,t})$ denote the state the individual is in at time t and given state variable $\varphi_{i,t}$. If $\xi_{i,t}(\varphi_{i,t}) = 1$ if the individual is in the high state; if $\xi_{i,t}(\varphi_{i,t}) = 0$ the individual is the low state.

Denote the full program of individual i's state-contingent choices:

$$c_i = \{c_{i,t}(\varphi_{i,t})\}_{t\in[0,1],\varphi_{i,t}\in\varphi},$$

$$x_i = \{x_{i,t}(\varphi_{i,t})\}_{t\in[0,1],\varphi_{i,t}\in\varphi}.$$

When in the high state the individual earns $y_i - \mathcal{T}(y_i)$ and when in the low state they receive benefits $\mathcal{B}(y_i)$. The individual can also earn additional income $f_{i,t}(x_{i,t}(\varphi_{i,t}))$. They face the flow budget constraint:

$$\dot{A}_{i,t}(\varphi_{i,t}) = \xi_{i,t}(\varphi_{i,t})\left(y_i - \mathcal{T}(y_i)\right) + (1 - \xi_{i,t}(\varphi_{i,t}))\mathcal{B}(y_i) + f_{i,t}(x_{i,t}(\varphi_{i,t})) - c_{i,t}(\varphi_{i,t})$$

with terminal condition: $A_{i1}(\varphi_{i,1}) > \bar{A}_i$ for all $\varphi_{i,1}$. They also face N additional constraints in each state $\varphi_{i,t}$ at each time t:

$$g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t}) \geq 0$$

The individual's problem is to choose the program (c_i, x_i) to solve:

$$\begin{aligned}
\max \int_t \int_{\varphi_{i,t}} & \left(\xi_{i,t}(\varphi_{i,t}) u^h_i(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t})) + (1 - \xi_{i,t}(\varphi_{i,t})) u^l_i(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}))\right) dF_{i,t}(\varphi_{i,t}) dt \\
& + \int_t \int_{\varphi_{i,t}} \lambda^A_{it}(\varphi_{i,t}) \left(\xi_{i,t}(\varphi_{i,t})\left(y_i - \mathcal{T}(y_i)\right) + (1 - \xi_{i,t}(\varphi_{i,t}))\mathcal{B}(y_i) + f_{i,t}(x_{i,t}(\varphi_{i,t})) - c_{i,t}(\varphi_{i,t})\right) dF_{i,t}(\varphi_{i,t}) dt \\
& + \int_{\varphi_{i,1}} \lambda^A_{i,1}(\varphi_{i,1})(A_{i1}(\varphi_{i,1}) - \bar{A}_i) dF_{i,1}(\varphi_{i,1}) \\
& + \sum_{n=1}^{N} \int_t \int_{\varphi_{i,t}} \lambda^n_{it}(\varphi_{i,t}) g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t}) dF_{i,t}(\varphi_{i,t}) dt
\end{aligned}$$

Denote the maximum function of this problem $V_i(\theta)$, written as a function of the policy parameters θ, which parameterizes modifications to the tax and benefit schedules. We assume the following regularity conditions:

ASSUMPTION 1 (Regularity Conditions). *Assume*

i. *Total lifetime utility is smooth, increasing and strictly quasi-concave in* (c, x)

ii. *The choices* (c, x) *that satisfy the constraints are convex*

iii. $V_i(\theta)$ *is differentiable*

where (i) and (ii) ensure the individual's problem has a unique solution and (iii) ensures the envelope theorem applies.

Consider the reform, $d\theta$, to the benefit and tax schedules parameterized by:

$$\begin{aligned}
\frac{d\mathcal{B}(y_i)}{d\theta} &= \phi_{\mathcal{B}}(y_i) \\
\frac{d\mathcal{T}(y_i)}{d\theta} &= \phi_{\mathcal{T}}(y_i) \times \frac{\int_{i'} \int_t \int_{\varphi_{i',t}} \phi_{\mathcal{B}}(y_{i'}) \left(1 - \xi_{i',t}(\varphi_{i',t})\right) dF_{i',t}(\varphi_{i',t}) dt di'}{\int_{i'} \int_t \int_{\varphi_{i',t}} \phi_{\mathcal{T}}(y_{i'}) \xi_{i',t}(\varphi_{i',t}) dF_{i',t}(\varphi_{i',t}) dt di'}
\end{aligned}$$

This reform is budget-neutral in the absence of any behavioral responses.

Assume that the constraints $g^n_{i,t}$ satisfy the regularity conditions:

$$\begin{aligned}
\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial \mathcal{B}(y_i)} &= -(1-\xi_{i,t}(\varphi_{i,t}))\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial c_{i,t}(\varphi_{i,t})}\\
\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial \mathcal{T}(y_i)} &= \xi_{i,t}(\varphi_{i,t})\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial c_{i,t}(\varphi_{i,t})}\\
\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial c_{i,s}(\varphi_{i,s})} &= 0 \quad \text{if } t \neq s
\end{aligned}$$

for $n = 1, \ldots, N$. See Chetty (2006) for a demonstration of the mildness of these conditions and an example of when they do not hold.

The impact of the reform $d\theta$ on individual i's expected utility is:

$$\begin{aligned}
\frac{dV_i}{d\theta} = &-\int_t\int_{\varphi_{i,t}}\left(\xi_{i,t}(\varphi_{i,t})\lambda^A_{i,t}(\varphi_{i,t}) - \sum_{n=1}^N \lambda^n_{i,t}(\varphi_{i,t})\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial \mathcal{T}(y_i)}\right) dF_{i,t}(\varphi_{i,t})dt\frac{d\mathcal{T}(y_i)}{d\theta}\\
&+\int_t\int_{\varphi_{i,t}}\left((1-\xi_{i,t}(\varphi_{i,t}))\lambda^A_{i,t}(\varphi_{i,t}) + \sum_{n=1}^N \lambda^n_{i,t}(\varphi_{i,t})\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial \mathcal{B}(y_i)}\right) dF_{i,t}(\varphi_{i,t})dt\frac{d\mathcal{B}(y_i)}{d\theta}
\end{aligned}$$

Under the regularity assumptions on the constraints, individual i's optimal consumption choice satisfies:

$$\frac{\partial u^s_i(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})} = \lambda^A_{i,t}(\varphi_{i,t}) - \sum_{n=1}^N \lambda^n_{i,t}(\varphi_{i,t})\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial c_{i,t}(\varphi_{i,t})}$$

for all t and $\varphi_{i,t}$. The first two assumptions imply for all t and $\varphi_{i,t}$:

$$\begin{aligned}
\sum_{n=1}^N \lambda^n_{i,t}(\varphi_{i,t})\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial \mathcal{T}(y_i)} &= \sum_{n=1}^N \lambda^n_{i,t}(\varphi_{i,t})\xi_{i,t}(\varphi_{i,t})\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial c_{i,t}(\varphi_{i,t})}\\
\sum_{n=1}^N \lambda^n_{i,t}(\varphi_{i,t})\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial \mathcal{B}(y_i)} &= -\sum_{n=1}^N \lambda^n_{i,t}(\varphi_{i,t})(1-\xi_{i,t}(\varphi_{i,t}))\frac{\partial g^n_{i,t}(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}); \varphi_{i,t})}{\partial c_{i,t}(\varphi_{i,t})}
\end{aligned}$$

Hence, we can re-write $dV_i/d\theta$:

$$\begin{aligned}
\frac{dV_i}{d\theta} = &-\int_t\int_{\varphi_{i,t}}\phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t})\frac{\partial u^h_i(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})}dF_{i,t}(\varphi_{i,t})dt\\
&\times\frac{\int_{i'}\int_t\int_{\varphi_{i',t}}\phi_{\mathcal{B}}(y_{i'})\left(1-\xi_{i',t}(\varphi_{i',t})\right)dF_{i',t}(\varphi_{i',t})dtdi'}{\int_{i'}\int_t\int_{\varphi_{i',t}}\phi_{\mathcal{T}}(y_{i'})\xi_{i',t}(\varphi_{i',t})dF_{i',t}(\varphi_{i',t})dtdi'}\\
&+\int_t\int_{\varphi_{i,t}}\phi_{\mathcal{B}}(y_i)(1-\xi_{i,t}(\varphi_{i,t}))\frac{\partial u^l_i(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})}dF_{i,t}(\varphi_{i,t})dt
\end{aligned}$$

We define willingness-to-pay, scaling $\frac{dV_i}{d\theta}$ as follows:

$$\text{WTP}_i \equiv \frac{dV_i}{d\theta} \times \left[\frac{\frac{1}{\int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)(1-\xi_{i,t}(\varphi_{i,t}))dF_{i,t}(\varphi_{i,t})dt}}{\frac{\int_t \int_{\varphi_{i,t}} \phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t}) \frac{\partial u_i^h(c_{i,t}(\varphi_{i,t}),x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})} dF_{i,t}(\varphi_{i,t})dt}{\int_t \int_{\varphi_{i,t}} \phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t})dF_{i,t}(\varphi_{i,t})dt}} \right]$$

The first factor expresses the utility gains per \$1 increase in lifetime benefit payments and the second factor converts to money-metric terms by comparing to the utility value of an unfunded tax cut (also expressed per \$1 worth of lifetime tax reduction).

Define the individual's state-specific average marginal utility of consumption:

$$\widetilde{\mathbb{E}}_i^h \left[\frac{\partial u^h(c,x)}{\partial c} \right] = \int_t \int_{\varphi_{i,t}} \frac{\phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t})}{\int_{t'} \int_{\varphi_{i,t'}} \phi_{\mathcal{T}}(y_i)\xi_{i,t'}(\varphi_{i,t'})dF_{i,t'}(\varphi_{i,t'})dt'} \frac{\partial u_i^h(c_{i,t}(\varphi_{i,t}),x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})} dF_{i,t}(\varphi_{i,t})dt$$

$$\widetilde{\mathbb{E}}_i^l \left[\frac{\partial u^l(c,x)}{\partial c} \right] = \int_t \int_{\varphi_{i,t}} \frac{\phi_{\mathcal{B}}(y_i)(1-\xi_{i,t}(\varphi_{i,t}))}{\int_{t'} \int_{\varphi_{i,t'}} \phi_{\mathcal{B}}(y_i)(1-\xi_{i,t'}(\varphi_{i,t'}))dF_{i,t'}(\varphi_{i,t'})dt'} \frac{\partial u_i^l(c_{i,t}(\varphi_{i,t}),x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})} dF_{i,t}(\varphi_{i,t})dt$$

We can then express willingness-to-pay for the reform:

$$\begin{aligned}\text{WTP}_i =& \frac{\widetilde{\mathbb{E}}_i^l \left[\frac{\partial u^l(c,x)}{\partial c} \right] - \widetilde{\mathbb{E}}_i^h \left[\frac{\partial u^h(c,x)}{\partial c} \right]}{\widetilde{\mathbb{E}}_i^h \left[\frac{\partial u^h(c,x)}{\partial c} \right]} + \\ & 1 - \frac{\int_t \int_{\varphi_{i,t}} \phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t})dF_{i,t}(\varphi_{i,t})dt}{\int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)(1-\xi_{i,t}(\varphi_{i,t}))dF_{i,t}(\varphi_{i,t})dt} \Big/ \frac{\int_{i'} \int_t \int_{\varphi_{i',t}} \phi_{\mathcal{T}}(y_{i'})\xi_{i',t}(\varphi_{i',t})dF_{i',t}(\varphi_{i',t})dtdi'}{\int_{i'} \int_t \int_{\varphi_{i',t}} \phi_{\mathcal{B}}(y_{i'}) \left(1-\xi_{i',t}(\varphi_{i',t})\right) dF_{i',t}(\varphi_{i',t})dtdi'},\end{aligned}$$

where the first expression on the right-hand side captures the lifetime value of insurance and the second expression captures the lifetime value of program cross-subsidization.

A.2. State-dependent utility

Consider the simple framework outlined in Section 2. Suppose the worker has state-dependent utility, which we capture through the marginal utility shifter $\rho > 0$, defined such that $u_i^{h\prime}(c) = u_i'(c)$ and $u_i^{l\prime}(c) = \rho u_i'(c)$.

In this case, the individual's WTP is:

$$\text{WTP}_i = \left(\frac{\rho u_i'(c_i^l) - u_i'(c_i^h)}{u_i'(c_i^h)} \right) + \left(1 - \frac{e_i \phi_{\mathcal{T}}(y_i)}{(1-e_i)\phi_{\mathcal{B}}(y_i)} \Big/ \frac{\int_{i'} e_{i'} \phi_{\mathcal{T}}(y_{i'})di'}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})di'} \right). \qquad \text{(A.1)}$$

The state-dependence term enters the insurance component of WTP, by scaling marginal utility in the low state. In this case, the consumption-based implementation uses the approximation:

$$\frac{\rho u_i'(c_i^l) - u_i'(c_i^h)}{u_i'(c_i^h)} \approx (\rho - 1) + \rho\gamma \frac{\Delta c_i}{c_i^h}$$

The state-dependence term captures factors that cause the same level of consumption expenditure to yield different marginal utilities across states. For example, leisure time may vary by state and act as a complement to or substitute for consumption. There may also be state-specific expenditures (e.g., work-related costs) that do not directly generate utility (Browning and Crossley 2001). Additionally, the extent to which individuals combine consumption expenditures with home production may differ across states (Aguiar and Hurst 2005).

The state-dependence parameter can also capture situations in which individuals face state-specific consumption prices (see Campos and Reggio 2020). Suppose the consumption price in the high state is ρ times the price in the low state, and utility is otherwise state independent. In this case, WTP takes the same form as in equation (A.1). The consumption-based implementation also leads to the same state-dependence-adjusted formula. However, it is necessary to correct observed expenditure changes to account for the state-specific price difference. Let x_i^s denote state-specific expenditures, which are observed in the data. These are related to consumption through:

$$\frac{\Delta c_i}{c_i^h} = \frac{x_i^h/\rho - x_i^l}{x_i^h/\rho} = \rho\left(\frac{\Delta x_i}{x_i^h}\right) - (\rho - 1)$$

B. Description of PSID Data

We use the Panel Study of Income Dynamics to estimate the sufficient statistics that allow us to quantify the welfare effects of unemployment insurance expansion. We use data from 1999-2019, the biennial so-called "new" PSID, which includes high quality information on consumption expenditures and asset holdings. For our baseline sample we focus on non-immigrant households.[1] We include both single and married households. We do not select the sample on the basis of the reference person's gender.[2] Our sample has non-missing information on key demographics

[1]This is an additional sample frame beginning from the late 1990s. We do, however, include the Survey of Economic Opportunity households which oversamples low income families. Results using survey weights are similar.

[2]Previously, the PSID referred to the reference person as the household head.

(age, education, and state of residence). We focus only on a sample of those where the reference person is aged between 23 and 60 and is participating in the labor force. Thus, we exclude observations where the reference person is retired, permanently disabled and neither working or looking for work, on sick (or maternity) leave or temporarily laid off, as well as those in education, who are a homemaker or in prison. In addition, to proxy for UI program eligibility we exclude those who never report employment, as well as periods immediately following a period out of the labor force. We include in our sample periods when the reference person is either in employment or unemployment, and we do not condition our sample on the employment status of their spouse. This is because we are explicitly interested in the employment risk facing households, including any self-insurance that may occur through spousal labor supply (Blundell et al. 2016).

To reduce the influence of measurement error, we also drop observations with extremely high asset values, as well as observations that exhibit large fluctuations in key outcomes of interest. Following, Blundell et al. (2016) we exclude households with net worth greater than $20 million to limit the impact of outliers and drop observations reporting extreme jumps (the bottom and top 0.25^{th} percentile) in wages, income and consumption to limit the role of measurement error. We do not use data displaying extreme "jumps" from one period to the next as we view this as most likely due to measurement error. A "jump" is defined as an extremely positive (negative) change from $t-2$ to t, followed by an extreme negative (positive) change from t to $t+2$. Formally, for each variable x, we construct the biennial log difference $\Delta_2 \ln(x_t)$, and drop the relevant variables for observation in the bottom 0.25 percent of the product $\Delta_2 \ln(x_t)\Delta_2 \ln(x_{t-2})$. Furthermore, in our analysis of permanent income we do not use earnings and wage data when the implied hourly wage is below one-half of the state minimum wage. Table B.1 provides summary statistics for our sample of interest.

To measure education we create five categories based on completed years of education: those completing less than high school, those who completed high school, those who complete high school and some college (including those who dropout of four year degrees, and those who attain a community college degree or other diploma), those who complete four years of college, and, finally, those who complete further study. Our measure of race is derived from the self-reported race of the household reference person, from which we define three race categories (White, Black, and other).

The majority of our empirical analysis focuses on the relationship between employment risk and both levels and growth rates in consumption. The PSID measures disaggregated consumption across a number of different categories of household expenditure which are designed to cover approximately 70% of aggregate consumption. Respondents can also indicate the period the expenditure covers. We convert all expenditures to the annual level (e.g., we multiply weekly expenditures by 52 and monthly expenditures by 12) and treat missing values in the consumption subcategories as zeros. We focus on expenditure categories that are measured consistently across survey waves. Our primary consumption measure of interest captures a range of utility-relevant expenditures comprising non-durable purchases plus service. We show, however, robustness of our empirical results to a variety of alternative consumption measures (including food consumption as in Gruber (1997) and Hendren (2017)) which we describe here. Consumption in the PSID is measured at the household level, to account for differences in household size we equivalize using the square root of the household size.

Baseline: Non-Durable Expenditure including Services. Our baseline measure includes a broad range of non-durable expenditures and services, as long as those services do not include an investment or durable component (for example, vehicle maintenance). To build our baseline consumption series we first construct a food expenditure series by summing food purchased to be consumed at home, food purchased to be consumed away from home, and those food purchases covered by the Supplemental Nutrition Assistance Program. Inclusion of food expenditures covered by the Supplemental Nutrition Assistance Program (formerly known as the Food Stamp Program and colloquially known as food stamps) is important because they are a relevant source of financing expenditures for low-income households, while Low and Pistaferri (2015) show that food stamps can act as substitutes for social insurance.

We then construct a household expenditure series for services without a durable component. We sum spending on home and auto-mobile insurance, utilities, parking costs and other direct transportation costs (such as bus fare and payments to taxis) that do not correspond to maintenance for a vehicle, as well as child care costs.

Finally, we combine the aggregated food expenditure series with household spending on gasoline expenses and the household expenditure series on services without a durable component.

Food Expenditure. To construct a series for food expenditure, which excludes the other components of our baseline measure, we sum food purchased to be consumed at home, food purchased to be consumed away from home, and those food purchases covered by food the Supplemental Nutrition Assistance Program.

Services Expenditure. To construct a series for broad services, inclusive of those that may relate to durables or have an investment component, we combine the services measured in our baseline expenditure with a set of additional spending categories. These additional categories include healthcare related spending (out-of-pocket payments including for hospital and nursing home stays, doctor visits, and prescription drugs as well as insurance premia), vehicle repairs, and payments for educational services or schooling costs such as school tuition.

Total (Non-Housing) Expenditure. We combine our baseline measure with the additional services described in the preceding paragraph (Services Expenditure) to produce a household-level series for total non-housing expenditures. As the PSID consumption categories are not designed to have full coverage, the name total expenditure is a misnomer. Furthermore, we continue to exclude the purchase of durables, such as vehicles, and memory goods (Hai et al. 2020), such as vacations, that have durable-like properties.

Total (Including Housing) Expenditure. Our final consumption measure incorporates a measure of housing services. We sum the total non-housing expenditures with the consumption value of housing services. For renters, we use reported rental expenditures. For homeowners, we approximate the rental equivalent flow of housing services as a 6 percent yield on the house price (Poterba and Sinai 2008).

We report summary statistics for our sample in Table B.1.

B.1. Constructing a Measure of Permanent Income

We include a measure of worker's permanent income in our statistical model for employment risk. Here we describe how we construct that measure.

We define permanent income as the predicted time-invariant individual component from the following log wage regression

$$\ln w_{i,t} = \beta_0 + \beta_1 age_{i,t} + \beta_2 age_{i,t}^2 + \beta_3 age_{i,t}^3 + \mu_i + \eta_t + u_{i,t},$$

where the dependent variable is the annual employment earnings of the reference person. We control for a third-order polynomial in age, and year fixed effects. We estimate this regression using only observations in which the worker is employed.

We then normalize the estimated individual fixed effects by their

	All	Employed	Unemployed
Ref. Person's Age (years)	39.57	39.69	37.56
	(10.31)	(10.29)	(10.31)
Married	0.53	0.55	0.30
	(0.50)	(0.50)	(0.46)
Number of Children	1.03	1.02	1.15
	(1.21)	(1.20)	(1.37)
Share White	0.55	0.56	0.32
	(0.50)	(0.50)	(0.47)
Share Black	0.13	0.13	0.13
	(0.33)	(0.33)	(0.33)
Ref. Person's Schooling (years)	13.78	13.84	12.76
	(2.45)	(2.44)	(2.41)
Unemployed	0.06	0.00	1.00
	(0.24)	(0.00)	(0.00)
Non-Durable Expenditure	20494.70	20852.78	14868.23
	(10760.47)	(10769.96)	(8,884.20)
Food Expenditure	9,596.98	9,741.15	7,331.58
	(5,750.15)	(5,771.84)	(4,866.71)
Services Expenditure	16226.51	16584.40	10603.04
	(13272.99)	(13344.30)	(10618.33)
Total Expenditure (Non-Housing)	25719.53	26211.44	17990.13
	(16316.70)	(16356.78)	(13496.92)
Total Expenditure (Incl. Housing)	38476.78	39289.33	25709.24
	(24918.80)	(24984.72)	(19899.40)
N	55671	52340	3331
Waves	3.65	3.45	1.51
Unique Households	15270	15188	2211

TABLE B.1. Summary Statistics for Our Sample of Interest
Computed from the 1999-2019 waves of the PSID. Tables shows and means and standard deviations in parenthesis. We denominate all dollar values using 2019 prices.

standard-deviation to construct our measure of permanent income:

$$\overline{Y}_i = \frac{\mu_i}{\sigma_\mu}.$$

This normalization provides an interpretable scale for permanent income. We use the resulting z-score in our estimation of employment risk, so the coefficient on permanent income captures the effect of a one standard deviation increase in permanent income.

C. Maximum Likelihood Estimator for Latent Heterogeneity

In the main text we specify group-specific heterogeneity as (equation (7) above):

$$\Delta_{i,t}^{FD} = \sum_{k \in K} \mathbb{1}\left[k(i) = k\right] \left(\delta_0^k + \delta_1^k U_{i,t}\right) + \beta X_{i,t} + \varepsilon_{i,t}, \tag{C.1}$$

and assume that $\varepsilon_{i,t} \sim N(0, \sigma_{g(i)}^2)$. In other words, we make a parametric restriction on the latent group specific density of errors. As discussed by Lewis et al. (2024), under this parametric restriction, identification of latent heterogeneity in consumption growth and the consumption response to the onset of unemployment exploits two complementarity sources of information: i) panel data information on persistent differences in households' consumption growth over time, and (ii) cross-sectional restrictions on the errors (which lack first-order autoregressive group structure under the Gaussian assumption) that can be distinguished from group-specific variation in the conditional mean. It is because we have relatively short panels for each household that lead us to leverage both sources of identification.

Let D collect the vector of individual-specific indicators defined by the partition $k(i)$ and $\mathbb{X}$ collect the covariates and unemployment indicators. Then, were the assignment to groups known with certainty, the complete-data likelihood is given by

$$L(\Delta^{FD}, \mathbb{X}, D; \chi) = \prod_{i=1}^{N} \prod_{t=1}^{T} \prod_{k=1}^{K} \left(\pi^k\right)^{d_{i,k} \times o_{i,t}} f\left(\Delta_{i,t}^{FD} | \underbrace{\delta_0^k + \delta_1^k U_{i,t} + \beta X_{i,t}}_{\mu^k(U_{i,t}, X_{i,t})}, (\sigma^g)^2\right)^{d_{i,k} \times o_{i,t}} \tag{C.2}$$

where $o_{i,t}$ is an indicator for whether the household is observed at time t, $f(\cdot|\mu, \sigma^2)$ is the density for the normal distribution with mean μ and variance σ^2. χ collects the parameters of interest: the unconditional probability of belonging to each group ($\pi^k = E[d_{i,k}]$), the (group-specific) parameters in the linear regression (δ^k and β), and the variance (σ^k).

As the assignment is not observed by the econometrician ex ante, we maximize the expected log-likelihood instead:

$$E_{D|\Delta^{FD}, \mathbb{X}}[\ln L(\Delta^{FD}, \mathbb{X}, D; \chi)] = \sum_{i=1}^{N} \sum_{t=1}^{T} o_{i,t} \sum_{g=1}^{G} \pi_i^g \left(\ln(\pi^k) + \ln f\left(\Delta_{i,t}^{FD} | \mu^k(U_{i,t}, X_{i,t}), (\sigma^k)^2\right)\right), \tag{C.3}$$

where

$$\pi_i^k = Pr(d_{i,k} = 1|\Delta_i^{FD}, \mathbb{X}_i) = \frac{\pi^k \prod_t f\left(\Delta_{i,t}^{FD}|\mu^g(U_{i,t}, X_{i,t}), (\sigma^k)^2\right)}{\sum_{k'=1}^K \pi^{k'} \prod_t f\left(\Delta_{i,t}^{FD}|\mu^h(U_{i,t}, X_{i,t}), (\sigma^{k'})^2\right)} \tag{C.4}$$

are posterior weights which capture the econometrician's ex post uncertainty over group assignment. We do not explicitly include other covariates in the conditioning set, instead these are valid posteriors conditional on the outcome and the regressors included in the linear model. We use the value of this posterior probability at our estimated parameters to construct the analogue $\hat{\pi}_i^k$ which we use to simulate the value of our sufficient statistics within sample with ${}^{\Delta c_i}/{}_{c_i}$ as $\sum_{k\in K} \hat{\pi}_i^k \hat{\delta}_1^k$.

As Lewis et al. (2024) highlight, the likelihood in (C.3) admits a sequential estimation procedure using the Expectation-Maximization (E-M) algorithm Dempster et al. (1977). We implement this estimation procedure using the R package `flexmix`. Due to the local convergence properties of the E-M algorithm we initialize the algorithm from 2000 different starting values to account for the possibility of local optima and select the estimates providing the largest value of the likelihood.

Inference on these objects is based on the Fisher Information Matrix.

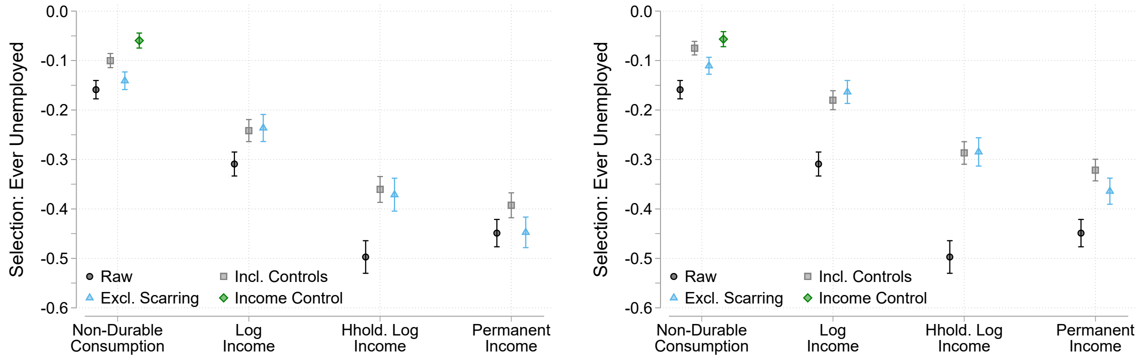
D. Additional Empirical Results

D.1. Robustness of Negative Selection Estimates to Alternative Controls

In our baseline empirical test of negative selection, we condition on a limited set of household characteristics. We exclude education and race controls, as redistributive policies are typically not designed with the explicit purpose of solely redistributing within these groups. Here, we assess the robustness of our findings to the inclusion of additional controls. Figure D.1 shows that average consumption differences remain robust when these controls are added. The right panel, which focuses on *within-group* differences, shows some attenuated effects of negative selection, reflecting the fact that we are controlling for other household characteristics relevant to consumption and labor market outcomes. Nevertheless, the quantitative impact of including these additional controls is modest.

D.2. Robustness of Consumption Decline Estimates to Alternative Consumption Measures

In addition to measuring consumption using our baseline expenditure measures, we also consider robustness to alternative consumption series. Table D.1 shows that



(A) Baseline

(B) Additional Controls

FIGURE D.1. Negative Selection Among the Unemployed

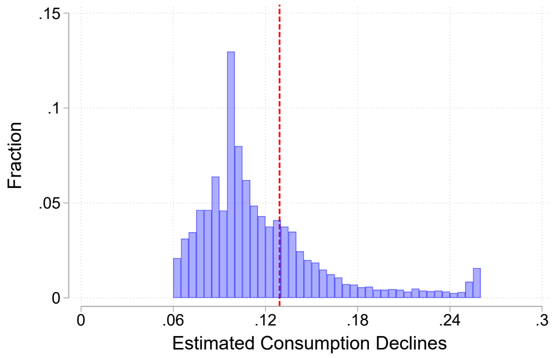
Notes: Authors' calculation from the PSID data. All regressions drop periods of unemployment. The no scarring sample conditions on periods before first observed unemployment. Raw specification includes no controls. In the left panel, all other specifications additionally control for a cubic function of age, indicator variables for whether they are married, a set of indicators for family size, and a set of year specific indicators. In the right panel we additionally control for gender, and a set of indicator variables capturing education status and race.

estimates of the average consumption decline are similar across these alternative measure and Figure D.2 shows how the distribution of consumption declines differs between our baseline measure and a measure of food expenditure as well as reporting the confidence interval for our estimate of the distribution.

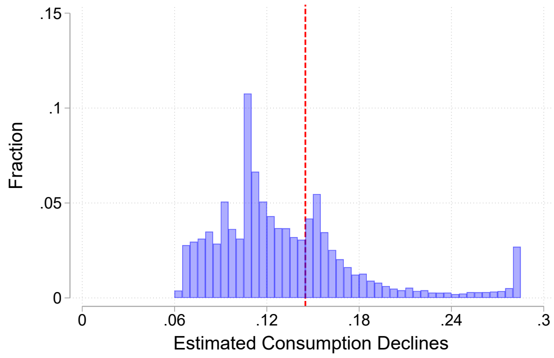
	Baseline	Alternative Measures			
	Non-durable	Food Expenditure	Services	Total (Non-Housing)	Total (Incl. Housing)
$\mathbb{1}\left[U_{i,t} = 1\right]$	-0.129^{***}	-0.145^{***}	-0.143^{***}	-0.129^{***}	-0.150^{***}
	(0.012)	(0.017)	(0.018)	(0.014)	(0.014)
Controls					
Age	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Year	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Family Size	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
N	37, 260	36, 997	37, 246	37, 289	37, 295

TABLE D.1. Consumption Declines by Alternative Consumption Measures ($\Delta c/c$)

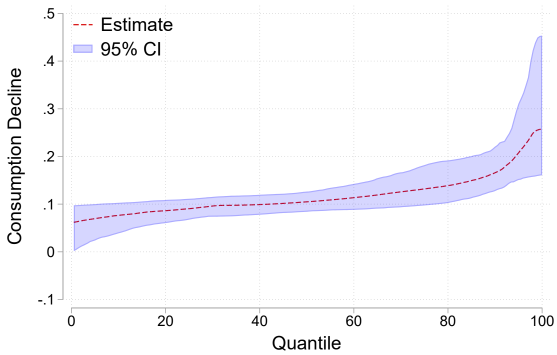
* $p < 0.10$, ** $p < 0.5$, *** $p < 0.01$. The first column reports the average consumption decline estimated (under the assumption of homogeneous consumption drops) for our baseline measure of consumption. The remaining columns report the consumption declines for alternative measures of consumption constructed from reported expenditures in the PSID data. Standard errors are clustered at the household-level. The number of observations across regressions differs due to item non-response.



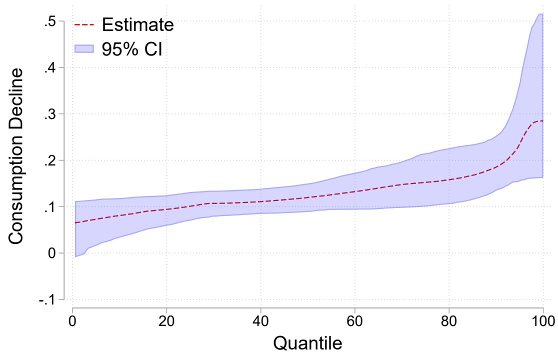
(A) Baseline (Non-Durable)



(B) Food Expenditures



(C) Baseline (Non-Durable)



(D) Food Expenditures

FIGURE D.2. Distribution of Estimated Consumption Declines for Alternative Consumption Series

Notes: Authors' estimates from the PSID data. The left panel reports our baseline estimates for our broad measure of non-durable consumption and the right panel reports results based on an alternative measure of consumption, food expenditures. The histogram (light blue bars) plots the individual proportional consumption declines constructed using the Gaussian mixture linear Regression-estimated parameters and individuals' posterior probabilities for each group, $\hat{\pi}_i^k$. For each household we compute the posterior-weighted effect of unemployment across the discrete group-specific unemployment consumption declines. The sample is defined as in the text. The Bayesian Information Criterion selects K = 3. The homogeneous estimate (red dashed vertical line) is estimated imposing K = 1 in our baseline specification. The second row shows our point estimates of each quantile along with the 95% confidence interval constructed from our bootstrap replications.

D.3. Observable Heterogeneity in Consumption Declines

Our approach to estimating unemployment-induced consumption declines avoids the need to specify observable determinants of its heterogeneity ex ante. We can, however, correlate our predicted declines with observables ex post. In Table D.2 we report the results of this exercise in order to understand observable determinants of heterogeneity in consumption declines.

We focus on proxies of a household's capacity to self-insure measured in periods of employment. This enables us to assess evidence for whether our estimated consumption declines are higher for those less able to self-insure.[3] To proxy for the availability of private savings, by which households may smooth consumption, we include indicators for each quartile of the distribution of liquid wealth and an indicator for whether or not the household owns their home. To construct liquid wealth, we follow Carroll and Samwick (1997) and sum the value of cash (including checking accounts) and savings, stocks owned outside of a retirement account, and bonds and government treasuries. These are the most liquid assets and can be used to smooth shocks in the short run. We additionally include indicators for quartiles of earned income, which captures both earning capacity and the size of the income loss a household will experience in unemployment. Finally, we include an indicator for marriage and cohabitation, which proxies for the additional insurance provided by the added work effect or spousal insurance.

We report the coefficient estimates from a regression of consumption declines (column 1) as well as an estimate of the conditional heteroskedasticity (column 2), which captures the extent to which the variability of consumption declines is impacted by observables. We find that, conditional on earning quartile, households with a larger ability to self insure (i.e., higher liquid assets, cohabiting, and home owning) have smaller consumption declines. Moreover, these factors also act to lower the dispersion of these consumption declines. These effects are economically and statistically significant.

We view this as both reinforcing the mechanisms highlighted by the existing literature, and highlighting the key role played by heterogeneity. Yet, we also find that these measures explain only a relatively small fraction of the variation in outcomes: an R^2 of 0.11. There are two interpretations of this finding. First, there is considerable latent heterogeneity that drives differences in the consumption exposure to unemployment (e.g., heterogeneous beliefs, preferences, or stochastic processes that lead to the optimality of different consumption behavior). Second, the proxies of a household's capacity to self insure are relatively weak (e.g., because liquidity is measured during employment and two years before job–loss). We conjecture that both are important, but note our approach is robust to both factors. This highlights a key advantage of the flexible approach we take to estimating the distribution of consumption declines.

[3] As Lewis et al. (2024) note in the context of marginal propensities to consume, a key advantage of using latent types to capture heterogeneity is that it implicitly allows for various observables to be included jointly, without loss of statistical power from the interaction of successively smaller groups in an interacted consumption decline specification.

	Consumption Decline	log(Variance)
Second Quartile Liquidity	-0.009^{***}	-0.253^{***}
	(0.001)	(0.044)
Third Quartile Liquidity	-0.013^{***}	-0.381^{***}
	(0.001)	(0.047)
Top Quartile Liquidity	-0.011^{***}	-0.211^{***}
	(0.001)	(0.051)
Second Quartile Earnings	-0.007^{***}	-0.171^{***}
	(0.001)	(0.046)
Third Quartile Earnings	-0.010^{***}	-0.234^{***}
	(0.001)	(0.050)
Top Quartile Earnings	-0.008^{***}	-0.158^{***}
	(0.001)	(0.056)
Home Owner	-0.013^{***}	-0.265^{***}
	(0.001)	(0.042)
Married	-0.011^{***}	-0.111^{**}
	(0.001)	(0.046)
N	47, 735	47, 735

TABLE D.2. Consumption Declines ($\Delta c/c$) and Proxies of Access to Self-Insurance
* $p < 0.10$, ** $p < 0.5$, *** $p < 0.01$. For each household we compute the posterior-weighted effect of unemployment across the discrete group-specific unemployment consumption declines. This is constructed using regression-estimated parameters and individuals' posterior probabilities for each group $\hat{\pi}_i^K$. We use the sample as defined as in the main text and consider periods of employment. The first column reports how individual consumption declines predicted by our Gaussian Linear Mixture Model vary with measures of ability to self-insure. The second column reports estimates of the conditional heteroskedasticity. As discussed above we define liquidity using the definition of "Very Liquid Assets" given by Carroll and Samwick (1997). Standard errors are clustered at the household level.

D.4. Modeling the Correlation between Employment Risk and Consumption Declines

In Table D.3 we report the underlying estimates that correspond to Figure 3

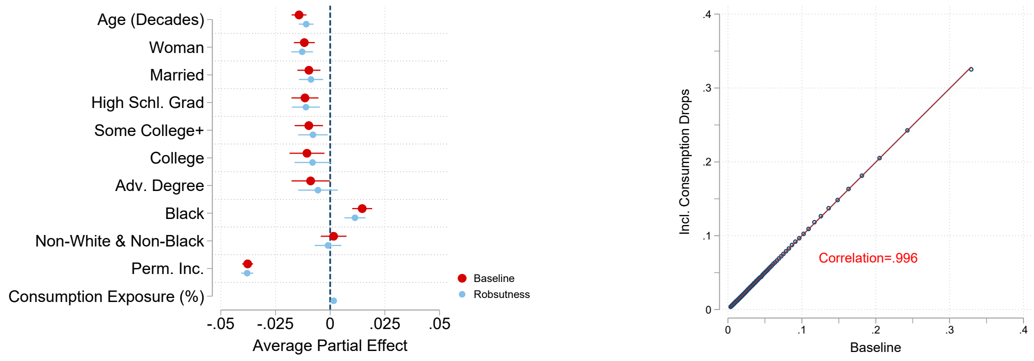
In our main empirical implementation we model employment risk as a function of observables. In this subsection we show that directly modeling the correlation between employment risk and our estimate of individual-specific consumption drops yields similar estimates of employment risk. To do so, we add to our model of employment risk (described in Section 4.2) a cubic function of the idiosyncratic consumption exposure.

Figure D.3 reports the results of this exercise. Panel (A) shows that the implied average partial effects in the extended model are very similar for all regressors included in the baseline model. These effects are slightly attenuated towards zero as

	Coefficient Estimates	Average Partial Effect
Age (Decades)	−2.013*	−0.014***
	(1.058)	(0.002)
Age^2	0.350	
	(0.273)	
Age^3	−0.021	
	(0.023)	
Woman	−0.284***	−0.012***
	(0.060)	(0.002)
High Schl. Grad	−0.257***	−0.012***
	(0.068)	(0.003)
Some College+	−0.214***	−0.010***
	(0.071)	(0.003)
College	−0.235**	−0.011***
	(0.092)	(0.004)
Adv. Degree	−0.194*	−0.009**
	(0.099)	(0.004)
Black	0.339***	0.015***
	(0.054)	(0.002)
Non-White & Non-Black	0.041	0.002
	(0.078)	(0.003)
Perm. Inc.	−1.048***	−0.038***
	(0.048)	(0.001)
PI^2	0.017	
	(0.014)	
PI^3	0.038***	
	(0.004)	
Constant	0.477	
	(1.315)	
N	52,996	52,996

TABLE D.3. Logit Estimation of e_i

Notes: Table reports the point estimates and standard errors for the coefficients of our estimated model of employment risk described in Section 4.2. In addition, we report the average partial effects and their corresponding standard errors which we display in Figure 3 in the main text.



(A) Average Partial Effects (B) Individual Predictions

FIGURE D.3. Correlation Between Consumption Exposure and Employment Risk

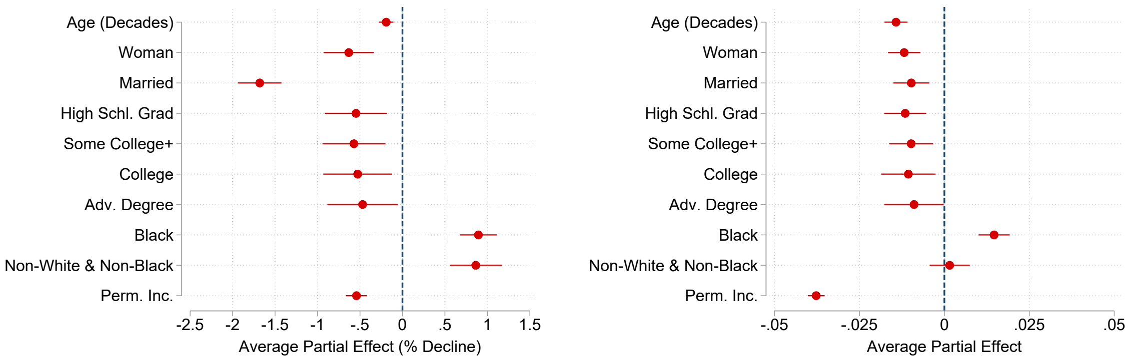
Notes: Authors' calculation from the PSID data. Our baseline measures of risk and consumption declines are estimated as we describe above and correspond to the model results summarized in Figures 2 and 3. Our robustness check includes a cubic function of the idiosyncratic consumption exposure in our model of employment risk. We use binned scatter plots with 100 bins in the right panel.

the additional measure of consumption exposure absorbs some of their explanatory power. However, the changes in effect sizes are economically insignificant. The effect of consumption exposure is statistically significant, but small. On average, an increase in the consumption exposure of 10 percentage points is associated with an increase in employment risk of 1.6 percentage points.

Panel (B) displays the average employment risk in our extended model as a function of our baseline estimates. There are no clear or systematic deviations in the predictions across models: values lie on or close to the 45-degree line and the predictions have a positive correlation of almost 1. Given we estimate the consumption declines in an alternative estimation procedure which complicates the problem of inference, we choose to use the simpler model of employment risk in our baseline analysis.

Figure 1 shows how both the estimated consumption declines and labor market risks vary with in-work consumption. Here we provide additional evidence on the link between employment risk and consumption declines. We regress our individual-level estimates of consumption declines on the same demographic and education information that we use to model employment risk. This directly illustrates the covariance between weights and consumption declines in our risk-weighted expectation.

Figure D.4 reports the average partial effects from this exercise (panel A), alongside our baseline estimates of average partial effects for employment risk (panel B, repeating Figure 3 in the main text). We find that factors that are systematically associated with smaller consumption declines are also associated with smaller probabilities of becoming unemployed and vice-versa. Consistent with the results shown above in Appendix D.3, we find a larger effect of marriage on consumption declines relative to employment risk, and a smaller permanent income gradient.



(A) Consumption Decline　　(B) Employment Risk

FIGURE D.4. Consumption Exposure, Employment Risk and Demographics
Notes: Authors' calculation from the PSID data. Our baseline measures of risk and consumption declines are estimated as we describe above and correspond to the model results summarized in Figures 3 and 2. We predict consumption declines as a function of the same variables we included in our estimated employment risk and report the average partial effects for both.

E. Willingness-to-Pay: Additional Results

E.1. Worker Preferences Across Alternative Reforms

Figure E.1 illustrates how WTP for UI expansions covaries across different reform types. We begin by dividing the sample into 20 equally sized groups based on WTP under a proportional-capped reform. Within each of these groups, we further divide individuals into 20 equally sized bins based on WTP for an alternative expansion. We then plot the joint distribution by showing average WTP across these bins for each reform. Panel (A) compares WTP under the baseline proportional-capped reform with WTP under a flat expansion. In the upper-right quadrant, we observe individuals that contribute relatively little to financing either reform. For these individuals WTP values converge to the 45-degree line, as, at this point, their contribution to reform financing is negligible and their reform valuations are therefore largely independent of its structure. Panel (B) shows a similar pattern when comparing the proportional-capped expansion with a proportional expansion.

By displaying the joint distribution, we can partition the WTP space into four mutually exclusive and exhaustive quadrants. The "Northeast" and "Southwest" quadrants contain individuals with consistent evaluation of both reforms: those in the Northeast derive positive surplus from either expansion, while those in the Southwest experience negative surplus under both. Two types of individuals fall in the Northeast quadrant: (i) those who in net terms contribute to funding the reform, but whose insurance value outweighs cross-subsidization effects, and (ii) those who benefit from cross-subsidization under both reforms—typically workers with high employment risk for whom the pooled reform price is below their actuarially fair price. Conversely, individuals in the Southwest quadrant find the pooled price more expensive than an actuarially fair individual price. This group disproportionately includes high-income individuals, who tend to face lower employment risk.

The differences across reform types highlight that workers may disagree about the desirability of UI expansions. Workers in the Southeast quadrant favor an expansion to the proportional-capped system but oppose a flat (Panel A) or proportional (Panel B) expansion. Conversely, those in the Northwest quadrant oppose a proportional-capped expansion, but favor a flat (Panel (A)) or proportional (panel (B)) expansion.

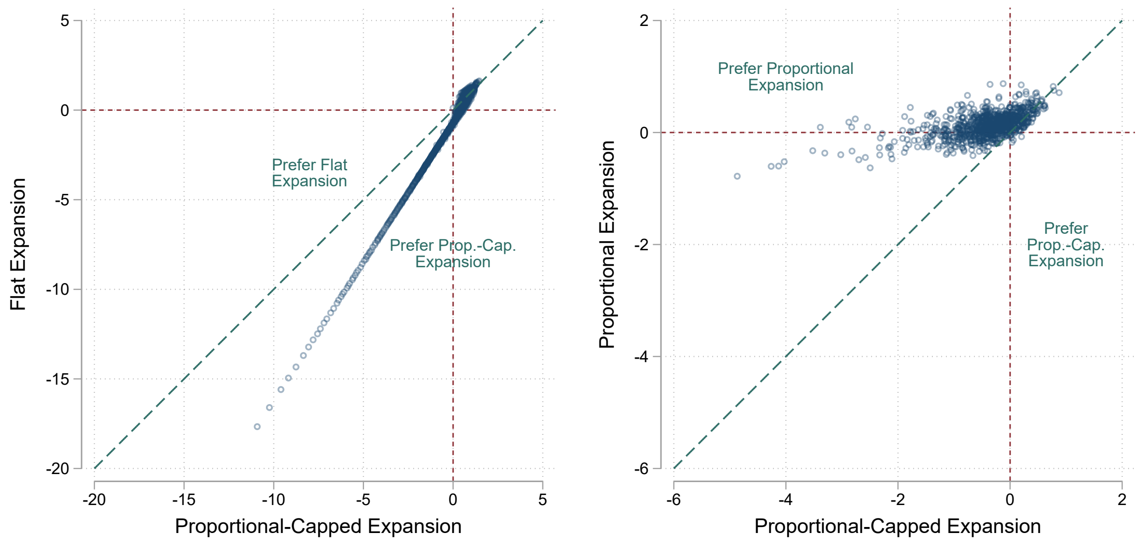
Additionally, we can classify individuals based on their relative ranking of alternative reforms. Those above the 45-degree line in Panel (A) prefer a flat expansion to a proportional-capped expansion. Similarly, those above the 45-degree line in panel (B) favor an increase in replacement rates to an expansion of the current proportional-capped system. Overall, 81% of workers prefer expanding the proportional-capped system to a uniform benefit increase, and 55% prefer it to an increase in the replacement rate.

E.2. Sensitivity to Preference Specification

Our willingness-to-pay results in Figure 4 are based on a coefficient of risk aversion of 3 and assume state-independent utility. Figure E.2 presents sensitivity analyses exploring the robustness of the implied WTP distributions to these assumptions. Since both risk-aversion and state dependence affect only the insurance component of WTP, we focus on that component.

Panel (A) reproduces the baseline distribution of the insurance value of WTP, conditional on household income (Figure 4(A)). Panel (B) introduces state-dependent utility: specifically, we assume that unemployed households face consumption prices 1.5% lower than those of employed households, following evidence from Kaplan and Menzio (2015) and Campos and Reggio (2020).[4] This assumption affects WTP in two offsetting ways (see Appendix A.2). First, lower prices raise the marginal value of a dollar when unemployed, increasing insurance value. Second, because

[4] Kaplan and Menzio (2015) find that married U.S. households with one unemployed member pay 1.5% lower prices; Campos and Reggio (2020) reports similar findings for Spanish households.



(A) Flat

(B) Proportional

FIGURE E.1. Individual Preferences over Reforms

Notes: Authors' estimates from the PSID data. We compare the proportional-capped system corresponding to current U.S. UI policy against a flat expansion (Panel (A)) and proportional (Panel (B)) reform.

observed expenditure understates true consumption for the unemployed, measured expenditure-based consumption drops overstate insurance values. Adjusting to correct the second effect dominates, modestly lowering the WTP distribution relative to the baseline.

Panels (C) and (D) show the WTP distribution assuming coefficients of relative risk aversions of 4 and 2, respectively. As expected, the distribution scales proportionately with risk aversion. However, because the distribution of WTP for realistic reforms is shaped primarily by cross-subsidization patterns (as shown in Figure 4), the overall distribution remains relatively insensitive to the calibrated level of risk aversion.

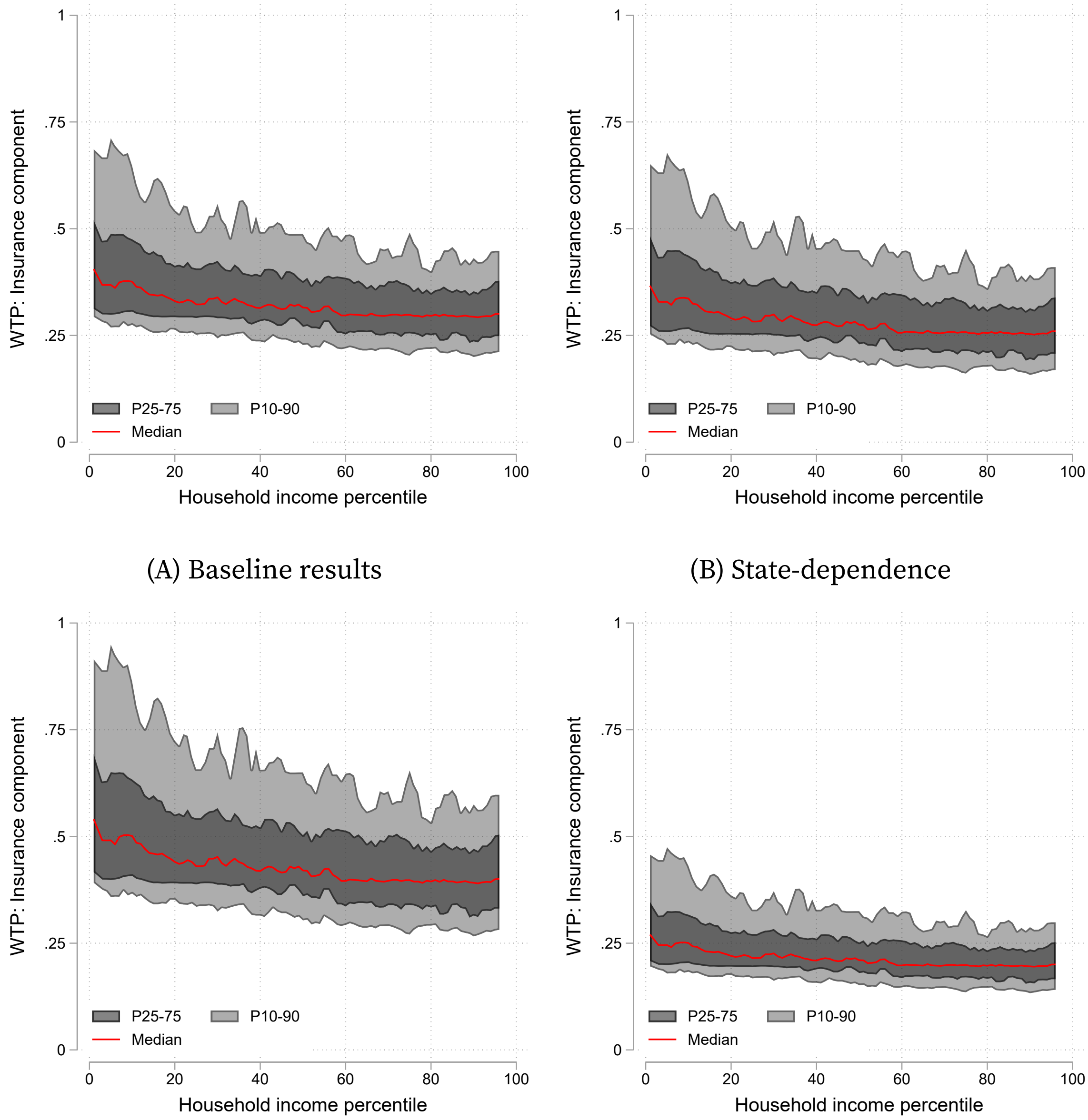
F. Normative Framework: Theory

F.1. Section 5.1 Derivations

Social welfare is given by

$$W(\theta) = \int_i \omega_i V_i(\theta) di$$

where individual expected utility, V_i, is defined in equation (1), and indexed by the reform parameter θ.



(A) Baseline results

(B) State-dependence

(C) High risk aversion

(D) Low risk aversion

FIGURE E.2. Willingness-to-Pay for Actuarially Fair UI Expansion by Income
Notes: Authors' estimates from the PSID data. Each panel summarizes the distribution of workers willingness-to-pay for an actuarially fair UI expansion (i.e,, the insurance component of overall WTP)—reporting the median, as well as 10th, 25th, 75th, 90th percentiles. We compute worker-level WTP using estimates of individual consumption declines. Panel (A) repeats our baseline results (Figure 4(A)). Panel (B) shows results assuming the price of consumption is 1.5% lower when unemployed. Panels (C) and (D) show results when the coefficient of relative risk aversion is 4 and 2, respectively. Quantiles are smoothed using a uniform rolling window of ± 1 percentiles.

We consider the reform parameterized by equation (2). Its impact on social welfare is:

$$\frac{dW}{d\theta} = \int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\omega_i u_i^{l\prime}(c_i^l)di - \int_i e_i\phi_{\mathcal{T}}(y_i)\omega_i u_i^{h\prime}(c_i^h)di\frac{d\tau}{db}$$

We focus on a budget-balanced reform, which implies that $\frac{d\tau}{db}$ is given by equation (9), repeated here:

$$\frac{d\tau}{db} = \frac{\int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di}{\int_i e_i\phi_{\mathcal{T}}(y_i)\,di}\left(1 + \frac{1}{\int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di}\left(\int_i \frac{d(1-e_i)}{d\theta}(\mathcal{T}(y_i) + \mathcal{B}(y_i))\,di\right)\right).$$

Note, we can re-express the fiscal externality (i.e., the portion of the tax adjustment needed to cover behavioral responses):

$$\begin{aligned}
\text{FE} \equiv & \frac{1}{\int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di}\left(\int_i \frac{d(1-e_i)}{d\theta}(\mathcal{T}(y_i) + \mathcal{B}(y_i))\,di\right) \\
= & \int_i \frac{(1-e_i)\phi_{\mathcal{B}}(y_i)}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})\,di'}\frac{1}{\phi_{\mathcal{B}}(y_i)}\frac{\mathcal{B}(y_i)}{1-e_i}\frac{d(1-e_i)}{d\theta}\frac{\mathcal{T}(y_i)+\mathcal{B}(y_i)}{\mathcal{B}(y_i)}di \\
= & \mathbb{E}^l\left[\epsilon^{(1-e,b)}\frac{\mathcal{T}(y)+\mathcal{B}(y)}{\mathcal{B}(y)}\right]
\end{aligned}$$

where the final line corresponds to equation (10), and uses the definition of the risk-weighted expectation (equation (8)) and the individual behavioral elasticity $\epsilon_i^{(1-e,b)} \equiv \frac{1}{\phi_{\mathcal{B}}(y_i)}\frac{d(1-e_i)}{db}\frac{\mathcal{B}(y_i)}{1-e_i}$. Note, in the definition of this elasticity db should be understood as the budget-balanced benefit change.

Using equations (8) and (9), we can then re-write $dW/d\theta$:

$$\frac{dW}{d\theta} = \int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di\left(\mathbb{E}^l\left[\omega u^{l\prime}(c^l)\right] - \mathbb{E}^h\left[\omega u^{h\prime}(c^h)\right] \times (1+\text{FE})\right)$$

Define the money-metric social welfare impact as:

$$\frac{dW^{MM}}{d\theta} = \frac{dW/d\theta}{\int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di}\Bigg/\frac{\partial W/\partial\tau}{\int_i e_i\phi_{\mathcal{T}}(y_i)\,di}$$

Division by $\int_i (1-e_i)\phi_{\mathcal{B}}(y_i)\,di$ expresses the change in social welfare per \$1 expansion in the benefit bill. The scaling by the inverse of $\frac{\partial W/\partial\tau}{\int_i e_i\phi_{\mathcal{T}}(y_i)\,di}$ expresses the social welfare change in money-metric terms, relative to the value of an unfunded tax cut. Note, $\frac{\partial W}{\partial\tau} = \int_i e_i\phi_{\mathcal{T}}(y_i)\,di\mathbb{E}^h\left[\omega_i u_i^{h\prime}(c_i^h)\right]$. As our implementation sets $\phi_{\mathcal{T}}(y_i) = 1$, this is equivalent to the effects of a lump-sum increase in high-state incomes.

Hence, we obtain equation (12), repeated here:

$$\frac{dW^{MM}}{d\theta} = \frac{\mathbb{E}^l[\omega u^{l\prime}(c^l)] - \mathbb{E}^h[\omega u^{h\prime}(c^h)]}{\mathbb{E}^h[\omega u^{h\prime}(c^h)]} - \text{FE}$$

Note that:

$$\begin{aligned}
\mathbb{E}^l[\omega u^{l\prime}(c^l)] - \mathbb{E}^h[\omega u^{h\prime}(c^h)] &= \mathbb{E}^l\left[\omega\left(u^{l\prime}(c^l) - \frac{e\phi_{\mathcal{T}}(y)}{(1-e)\phi_{\mathcal{B}}(y)} \Big/ \frac{\int_{i'} e_{i'}\phi_{\mathcal{T}}(y_{i'})di'}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})di'} u^{h\prime}(c^h)\right)\right] \\
&= \mathbb{E}^l\left[\omega u^{h\prime}(c^h)\left(\frac{u^{l\prime}(c^l) - u^{h\prime}(c^h)}{u^{h\prime}(c^h)} + \left(1 - \frac{e\phi_{\mathcal{T}}(y)}{(1-e)\phi_{\mathcal{B}}(y)} \Big/ \frac{\int_{i'} e_{i'}\phi_{\mathcal{T}}(y_{i'})di'}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})di'}\right)\right)\right] \\
&= \mathbb{E}^l[\omega u^{h\prime}(c^h)]\mathbb{E}^l\left[\lambda\left(\frac{u^{l\prime}(c^l) - u^{h\prime}(c^h)}{u^{h\prime}(c^h)} + \left(1 - \frac{e\phi_{\mathcal{T}}(y)}{(1-e)\phi_{\mathcal{B}}(y)} \Big/ \frac{\int_{i'} e_{i'}\phi_{\mathcal{T}}(y_{i'})di'}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})di'}\right)\right)\right],
\end{aligned}$$

where $\lambda_i \equiv \frac{\omega_i u_i^{h\prime}(c_i^h)}{\mathbb{E}^l[\omega u^{h\prime}(c^h)]}$. Hence we obtain the decomposition in equation (13),

$$\frac{dW^{MM}}{d\theta} = \bar{\lambda}\mathbb{E}^l\left[\lambda\left(\frac{u^{l\prime}(c^l) - u^{h\prime}(c^h)}{u^{h\prime}(c^h)} + 1 - \frac{e\phi_{\mathcal{T}}(y)}{(1-e)\phi_{\mathcal{B}}(y)} \Big/ \frac{\int_{i'} e_{i'}\phi_{\mathcal{T}}(y_{i'})di'}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})di'}\right)\right] - \text{FE},$$

where $\bar{\lambda} \equiv \frac{\mathbb{E}^l[\omega u^{h\prime}(c^h)]}{\mathbb{E}^h[\omega u^{h\prime}(c^h)]}$. Note that:

$$\mathbb{E}^l\left[\lambda\left(1 - \frac{e\phi_{\mathcal{T}}(y)}{(1-e)\phi_{\mathcal{B}}(y)} \Big/ \frac{\int_{i'} e_{i'}\phi_{\mathcal{T}}(y_{i'})di'}{\int_{i'}(1-e_{i'})\phi_{\mathcal{B}}(y_{i'})di'}\right)\right] = 1 - \mathbb{E}^h[\lambda] = 1 - \frac{1}{\bar{\lambda}}$$

hence we have the decomposition in equation (14):

$$\frac{dW^{MM}}{d\theta} = \bar{\lambda}\mathbb{E}^l\left[\lambda\left(\frac{u^{l\prime}(c^l) - u^{h\prime}(c^h)}{u^{h\prime}(c^h)}\right)\right] + (\bar{\lambda} - 1) - \text{FE}.$$

F.2. Extension to General Model

Under the general model outlined in Appendix A.1, the impact of the reform on social welfare is given by:

$$\begin{aligned}
\frac{dW}{d\theta} = &- \int_i \omega_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t}) \frac{\partial u_i^h(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})} dF_{i,t}(\varphi_{i,t})dtdi \times \frac{d\tau}{db} \\
&+ \int_i \omega_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)(1 - \xi_{i,t}(\varphi_{i,t})) \frac{\partial u_i^l(c_{i,t}(\varphi_{i,t}), x_{i,t}(\varphi_{i,t}))}{\partial c_{i,t}(\varphi_{i,t})} dF_{i,t}(\varphi_{i,t})dtdi,
\end{aligned}$$

where budget-balance requires:

$$\frac{d\tau}{db} = \frac{\int_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)\left(1-\xi_{i,t}(\varphi_{i,t})\right) dF_{i,t}(\varphi_{i,t})dtdi}{\int_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t})dF_{i,t}(\varphi_{i,t})dtdi} \times \left(1 + \frac{1}{\int_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)\left(1-\xi_{i,t}(\varphi_{i,t})\right) dF_{i,t}(\varphi_{i,t})dtdi} \times \right.$$

$$\left. \int_i \int_t \int_{\varphi_{i,t}} \frac{d(1-\xi_{i,t}(\varphi_{i,t}))}{d\theta}(\mathcal{T}(y_i) + \mathcal{B}(y_i))dF_{i,t}(\varphi_{i,t})dtdi \right)$$

$$= \frac{\int_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)\left(1-\xi_{i,t}(\varphi_{i,t})\right) dF_{i,t}(\varphi_{i,t})dtdi}{\int_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t})dF_{i,t}(\varphi_{i,t})dtdi} \times \left(1 + \underbrace{\mathbb{E}^l\left[\epsilon \times \frac{\mathcal{T}(y_i)+\mathcal{B}(y_i)}{\mathcal{B}(y_i)}\right]}_{\text{FE}}\right).$$

ϵ_i is given by:

$$\epsilon_i = \int_t \int_{\varphi_{i,t}} \frac{d(1-\xi_{i,t}(\varphi_{i,t}))}{db} dF_{i,t}(\varphi_{i,t})dt \frac{1}{\phi_{\mathcal{B}}(y_i)} \frac{\mathcal{B}(y_t)}{\int_t \int_{\varphi_{i,t}}(1-\xi_{i,t}(\varphi_{i,t}))dF_{i,t}(\varphi_{i,t})dt}$$

and the low-state risk-weighted expectation over any i-specific variable x_i is given by:

$$\mathbb{E}^l\left[x\right] \equiv \int_i \left(\frac{\int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)(1-\xi_{i,t}(\varphi_{i,t}))dF_{i,t}(\varphi_{i,t})dt}{\int_{i'} \int_t \int_{\varphi_{i',t}} \phi_{\mathcal{B}}(y_{i'})(1-\xi_{i',t}(\varphi_{i',t}))dF_{i',t}(\varphi_{i',t})dtdi'} \right) x_i di$$

with the high-state risk-weighted expectation defined analogously (see equation (8)).

Under the standard assumption that when a worker finds a job they remain employed thereafter, $D_i \equiv \int_t \int_{\varphi_{i,t}}(1-\xi_{i,t}(\varphi_{i,t}))dF_{i,t}(\varphi_{i,t})dt$ is the worker's expected unemployment duration, ϵ_i is the elasticity of worker i's unemployment duration with respect to the benefit change $\phi_i(y_i)db$ they face under the budget-balanced reform, and the low risk-weighted expectation weights by individuals' expected benefit rise, which depends on their expected duration and benefit adjustment, $\phi_i(y_i)$.

To express the social welfare impact per \$ of increased benefit expenditure and in money-metric terms, relative to an unfunded tax cut, we rescale $dW/d\theta$ according to:

$$\frac{dW^{MM}}{d\theta} = \frac{dW/d\theta}{\int_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{B}}(y_i)\left(1-\xi_{i,t}(\varphi_{i,t})\right) dF_{i,t}(\varphi_{i,t})dtdi} \Big/ \frac{\partial W/\partial\tau}{\int_i \int_t \int_{\varphi_{i,t}} \phi_{\mathcal{T}}(y_i)\xi_{i,t}(\varphi_{i,t})dF_{i,t}(\varphi_{i,t})dtdi}.$$

After a couple of lines of algebra analogous to those in the preceding subsection, we obtain:

$$\frac{dW^{MM}}{d\theta} = \frac{\mathbb{E}^l\left[\omega\frac{\partial u^l(c,x)}{\partial c}\right] - \mathbb{E}^h\left[\omega\frac{\partial u^h(c,x)}{\partial c}\right]}{\mathbb{E}^h\left[\omega\frac{\partial u^h(c,x)}{\partial c}\right]} - \text{FE}$$

F.3. Relationship with Marginal Value of Public Funds

Our normative analysis considers a reform to UI generosity that raises benefits, financed by a flat increase in payroll taxation. We focus on this combined reform because the structure of the U.S. UI system links UI benefits to a hypothecated tax. However, it is instructive to consider the benefit and tax components separately. Doing so clarifies the relationship between our analysis and the marginal value of public funds (MVPF) framework (see Hendren and Sprung-Keyser 2020).

The MVPF for an unfunded benefit rise is:

$$\text{MVPF}^{\mathcal{B}} = \frac{\mathbb{E}^l[\text{WTP}^{\mathcal{B}}]}{1+\text{FE}^{\mathcal{B}}}$$

where the willingness-to-pay is:

$$\text{WTP}_i^{\mathcal{B}} = \left(\frac{u_i^{l\prime}(c_i^l) - u_i^{h\prime}(c_i^h)}{u_i^{h\prime}(c_i^h)} + 1\right)$$

and $\text{FE}^{\mathcal{B}}$ denotes the associated fiscal externality from behavioral responses.

Willingness-to-pay to avoid a flat tax increase is 1, and the MVPF for the tax rise is:

$$\text{MVPF}^{\mathcal{T}} = \frac{1}{1+\text{FE}^{\mathcal{T}}},$$

where $\text{FE}^{\mathcal{T}}$ is the fiscal externality associated with the payroll tax increase.

To assess whether the combined reform is socially valuable, we require that the MVPF of the benefit expansion exceeds that of the tax increases, with each MVPF valued at the incidence-weighted average marginal social welfare weights of affected groups:

$$\mathbb{E}^l\left[g\left(\frac{\text{WTP}^{\mathcal{B}}}{\mathbb{E}^l[\text{WTP}^{\mathcal{B}}]}\right)\right]\frac{\mathbb{E}^l[\text{WTP}^{\mathcal{B}}]}{1+\text{FE}^{\mathcal{B}}} > \mathbb{E}^h[g]\frac{1}{1+\text{FE}^{\mathcal{T}}}$$

Rewriting this inequality, we obtain:

$$\frac{\mathbb{E}^l[g]}{\mathbb{E}^h[g]}\mathbb{E}^l\left[\frac{g}{\mathbb{E}^l[g]}\text{WTP}^{\mathcal{B}}\right]-\frac{1+\text{FE}^{\mathcal{B}}}{1+\text{FE}^{\mathcal{T}}}>0$$

$$\frac{\mathbb{E}^l[g]}{\mathbb{E}^h[g]}\mathbb{E}^l\left[\frac{g}{\mathbb{E}^l[g]}\left(\frac{u^{l\prime}(c^l)-u^{h\prime}(c^h)}{u^{h\prime}(c^h)}\right)\right]+\left(\frac{\mathbb{E}^l[g]}{\mathbb{E}^h[g]}-1\right)-\left(\frac{1+\text{FE}^{\mathcal{B}}}{1+\text{FE}^{\mathcal{T}}}-1\right)>0.$$

Recognizing that $\bar{\lambda}=\frac{\mathbb{E}^l[g]}{\mathbb{E}^h[g]}$, $\lambda_i=\frac{g_i}{\mathbb{E}^l[g]}$ and that the fiscal externality of the combined reform in the main text is equivalent to $\text{FE}=\frac{1+\text{FE}^{\mathcal{B}}}{1+\text{FE}^{\mathcal{T}}}-1$, makes clear that the left-hand side of the inequality is exactly to the welfare condition in equations (12) and (14).[5]

G. Normative Framework: Implementation

G.1. Measuring Fiscal Externality

Our fiscal externality term decomposes as:

$$\text{FE}=\mathbb{E}^l\left[\epsilon^{(1-e,b)}\right]+\mathbb{E}^l\left[\epsilon^{(1-e,b)}\frac{\mathcal{T}(y)}{\mathcal{B}(y)}\right], \tag{G.1}$$

where we refer to the first component as the direct fiscal externality from UI expansion, arising from higher benefit payments. The second component captures the indirect effect, which operates through changes in income tax payments.

The first component depends on risk weights (in the expectation), which we construct using our employment risk model and observed data on worker wages, along with the behavioral elasticity with respect to UI benefits. We draw on elasticity estimates from Chetty (2008), who reports how this elasticity varies across wealth quartiles (see Table 2, p. 204). We assign each household in our data the corresponding elasticity based on its wealth quartile. Following Chetty (2008), we adjust these elasticities to account for benefit exhaustion after six months. Specifically, we scale them by the ratio of average benefit and to average unemployment duration (15.8/24.3) (see also discussion in Schmieder and Von Wachter (2016)).

The indirect component of the fiscal externality depends on risk weights, behavioral elasticities, and, additionally, the state-specific tax position of workers. We compute these tax positions using the NBER TAXSIM (version 32) U.S. tax and transfer simulator.

[5] In practice, it is likely that $\text{FE}^{\mathcal{B}}\approx\text{FE}$ and $\text{FE}^{\mathcal{T}}\approx 0$, since the fiscal externality of the combined reform is likely primarily driven by the benefit expansion rather than the small funding-side payroll tax increase. However, the above equivalence does not rely on this being the case.

Specifically, for all observations in which the household's reference person is employed, we compute household annual tax liability, accounting for spousal earnings in married households (assuming joint filing) and household composition. This yields an estimate of the household's tax liability in the high state, denoted by T_i^h.

To construct the net tax liability in the low state, we assume the reference person experiences an unemployment spell of average duration and receives unemployment benefits equal to the lower of 50% of wages or the UI earnings cap (approximately 41% of average wages). Denote these UI payments by ui_i. We then compute the household's tax liability in the low state, denoted by T_i^l.

The empirical counterpart of $\frac{\mathcal{T}(y)}{\mathcal{B}(y)}$ is given by:

$$\frac{T_i^h - T_i^l}{ui_i},$$

which reflects the difference in state-specific income tax liabilities between employment and unemployment, scaled by unemployment benefit payments.

G.2. Average Implementation

Our full implementation of the flat reform entails evaluating the expression:

$$\frac{\mathbb{E}^l[g_y\tilde{g}]}{\mathbb{E}^h[g_y\tilde{g}]}\mathbb{E}^l\left[\frac{g_y\tilde{g}}{\mathbb{E}^l[g_y\tilde{g}]}\gamma\frac{\Delta c}{c^h}\right] + \left(\frac{\mathbb{E}^l[g_y\tilde{g}]}{\mathbb{E}^h[g_y\tilde{g}]} - 1\right) - \left(\mathbb{E}^l\left[\epsilon^{(1-e,b)}\right] + \mathbb{E}^l\left[\epsilon^{(1-e,b)}\frac{\mathcal{T}(y)}{\mathcal{B}(y)}\right]\right),$$

which is obtained by combining equations (4), (12), (14), (15) and (G.1). Recall that the welfare weights include a cross-income (inverse-optimum) component g_y and a within-income-group component $\tilde{g}_i = \frac{(c_i^h)^{-\gamma}}{\mathbb{E}[(c_i^h)^{-\gamma}|y]}$

In the average implementation, we treat the employed and unemployed populations as single individuals with average values, $\mathbb{E}^s[x]$ for $s \in \{h, l\}$, for the income-based welfare weight (g_y), consumption (c_i^h), consumption drop ($\Delta c/c_h$), benefit entitlement ($\mathcal{B}$), and tax liability ($\mathcal{T}$). This leads to the simplified expression:

$$\frac{\mathbb{E}^l[g_y]}{\mathbb{E}^h[g_y]} \times \gamma\mathbb{E}^l\left[\frac{\Delta c}{c^h}\right] + \left(\frac{\mathbb{E}^l[g_y]}{\mathbb{E}^h[g_y]} - 1\right) - \left(\mathbb{E}^l\left[\epsilon^{(1-e,b)}\right] + \mathbb{E}^l\left[\epsilon^{(1-e,b)}\right]\frac{\mathbb{E}^l\left[\mathcal{T}(y)\right]}{\mathbb{E}^l\left[\mathcal{B}(y)\right]}\right)$$

The average implementation captures differences in average welfare weights between the employed and unemployed populations, but it ignores within-group heterogeneity. As a result, this average implementation is invariant across reform types, since difference between reforms stem from variation in their incidence across the unemployed population.

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