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The Rise in Disability Recipiency and the Decline in Unemployment

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Working Paper 01-15
May 2001

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Abstract

Between 1984 and 2000, the share of non-elderly adults receiving benefits from the Social Security Disability Insurance (DI) and Supplemental Security Income (SSI) programs rose from 3.1 to 5.3 percent. We trace this growth to reduced screening stringency and, due to the interaction between growing wage inequality and a progressive benefits formula, a rising earnings replacement rate. We explore the implications of these changes for the level of labor force participation among the less skilled and their employment responses to adverse employment shocks. Following program liberalization in 1984, DI application and recipiency rates became two to three times as responsive to plausibly exogenous labor demand shocks. Contemporaneously, male and female high school dropouts became increasingly likely to exit the labor force rather than enter unemployment in the event of an adverse shock. The liberalization of the disability program appears to explain both facts. Accounting for the role of disability in inducing labor force exit among the low-skilled unemployed, we calculate that the U.S. unemployment rate would be two-thirds of a percentage point higher at present were it not for the liberalized disability system.

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Introduction

The federal disability programs are the largest social transfer programs in the United States directed towards non-elderly adults. Annual disability expenditures exceed that of welfare (TANF), Unemployment Insurance, and the Earned Income Taxed Credit combined. In the past two decades, there have been two important changes to these programs affecting eligibility and generosity. The first, occurring in 1984, liberalized the disability determination process, reversing a dramatic reduction in the disability beneficiary population that was underway. The second occurred more gradually. Because disability benefits are indexed to the mean wage in the economy, the widening dispersion of earnings throughout the 1980s and 1990s coupled with the progressivity of the disability benefits formula substantially raised the effective replacement rate – the ratio of disability income to prior earnings – for low skilled workers. Together, these changes appear to have induced unprecedented growth in disability recipiency. From 1984 to 2000, the number of non-elderly adults receiving benefits from the Social Security Disability Insurance (DI) and/or Supplemental Security Income (SSI) programs more than doubled from 3.8 to 7.7 million. At present, 5.3 percent of adults ages 25 to 64 receive either SSI, DI or both. Accompanying this growth, disability recipients became significantly younger and substantially more likely to suffer from comparatively low mortality maladies such as mental disorders and pain.

This paper assesses the contribution of program liberalization and rising replacement rates to the labor market outcomes of low skilled individuals. After describing the recent evolution of the disability program, we offer a simple dynamic programming model of the decision to apply for disability benefits as a function of earnings, health, labor market conditions, and program generosity. Our model demonstrates that because labor force non-participation is a de facto pre-condition for disability application, the generosity of disability benefits – both replacement rates and screening stringency –have two distinct effects on labor force participation among potential beneficiaries. The first is a 'level' effect. More liberal program screening and higher benefit levels directly reduce the level of labor force participation among potential beneficiaries. The second is an interaction effect. Because the unemployed face lower opportunity costs of applying for disability than the employed – in particular, they do not have to leave employment – there is a subset of workers who will *not* apply for disability while employed but will apply in the event of job loss. We refer to these individuals as 'conditional applicants.'

A primary implication of the model is that increasing program generosity and deteriorating labor market conditions – circumstances facing the low skilled during the period under study – serve to expand the relative size of this conditional group; and the sum of these two factors is more potent than their individual components. As the model demonstrates, increasing program generosity and worsening labor market conditions do *not* for the most part spur workers to exit employment directly; rather, they increase the option value of disability application. Consequently, in the event of job loss, a greater fraction of workers exercises this option by exiting the labor force to seek benefits. Hence, the secular changes in the disability program since 1984 are likely to have reduced the aggregate unemployment rate by spurring job losers to withdraw from the labor force.

We take these predictions to the data in two stages. First, we exploit cross-state variation in the administrative clampdown on disability awards over the period of 1979 to 1984 to study the effects of program stringency on the *level* of labor force participation among low skilled workers, particularly high school dropouts. Because the disability benefits formula is progressive but not indexed to regional wage levels, workers in low wage states face significantly higher earnings replacement rates than workers in high wage states. This gives rise to substantial cross-state variation in disability application and recipiency rates. As we demonstrate, these cross-sectional differences interacted significantly with national shifts in the supply of benefits during the retrenchment and reform periods. In particular, the share of state residents deterred or encouraged to apply for benefits in each era depended significantly upon effective state replacement rates.

By exploiting these cross-state shifts in the effective supply of benefits, we estimate the impact of the disability program on the *level* of labor force participation. During the retrenchment, the labor force participation of males with less than a high school degree *increased* most in states with the largest reductions in the supply of benefits, and vice versa during the subsequent liberalization. The estimates indicate that these exogenous supply shifts raised male high school dropout labor force participation by as much as 5 percentage points over 1979 – 1984, and lowered it by a slightly larger amount thereafter. Notably, no similar relationship between labor force participation and the supply of disability benefits is observed for other education groups.

We next explore the interaction of the disability program with adverse demand shocks. By exploiting plausibly exogenous cross-state variation in labor demand shocks, we examine whether the sensitivity of disability application and labor force exit to adverse demand shocks has increased as predicted by the model. Following disability reform, application and award rates for given demand shocks rose secularly, reaching two to three times their pre-reform levels by the 1990s. Paralleling this change was a substantial increase in the rate of labor force exit among displaced high school dropouts. During the pre-reform era of 1978 – 1984, approximately 40 percent of high school dropouts who lost their jobs due to adverse demand shocks transitioned into unemployment, while the remaining 60 percent exited the labor force. From 1984 forward, however, the share entering unemployment in response to job loss declined dramatically. Simple calculations suggest that increased disability application propensity is likely to account for this behavioral change.

To gauge the importance of this phenomenon, we calculate the demand contraction experienced by high school dropouts over 1984 – 1998 to form a counterfactual labor force participation figure net of disability. Our calculations suggest that the aggregate U.S. unemployment rate would at present be approximately 0.65 percentage points higher were it not for the liberalization of the disability program in 1984. This estimate will be biased upwards if part of the reduced sensitivity of unemployment to adverse demand shocks is due to other factors. Given that we consider the impact of the DI program only on the 12 percent of the population that is composed of high school dropouts, however, it seems plausible that our estimates may potentially underestimate the total effect.[1]

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[1] Our analysis is related to Parsons (1981), Bound (1989), Bound and Waidman (1992), Gruber and Kubik (1995), Aarts, Burkhauser, and De Jong (1996) and Gruber (2000), who study the impact of disability benefits on labor supply in Canada, the Netherlands, and the United States; and to Lewin-VHI, Inc. (1995), Rupp and Stapleton (1995), Black, Daniel and Sanders (1998) and Stapleton et al. (1998) who analyze the importance of the economic climate to disability application and recipiency. Bound and Burkhauser (1999) provide an excellent overview of the

1. Context: Thefederal disability programs

The federal government provides cash and medical benefits to individuals with disabilities through the Social Security Disability Insurance (DI) and Supplemental Security Income (SSI) programs. The medical eligibility criteria for the two programs are identical, requiring that an individual have a medically determinable impairment that prevents him or her from engaging in "substantial gainful work." SSI benefits are means-tested and do not require any prior work history while DI benefits are an increasing function of prior earnings and are not means-tested.[2] To apply for benefits, an individual must submit detailed medical, income, and asset information to a federal Social Security Administration (SSA) office, which makes the disability determination.

a. Clampdown and liberalization

Since the founding of the DI program in 1956 and the federalization of SSI in 1973, the number of individuals receiving disability benefits has risen substantially. During the late 1970s, a rapidly growing disability recipient pool coupled with declining termination rates raised concerns that many ineligible individuals were receiving disability benefits. In response, the Social Security Administration tightened medical eligibility criteria and exercised firmer control over the state boards that interpret SSA's eligibility standards. Acceptance rates for both programs fell substantially. The fraction of DI applicants awarded benefits fell from 45 percent in 1976 to 32 percent in 1980, accompanied by an even more pronounced 19 percentage point decline in SSI acceptance rates.

Augmenting this administrative action, Congress passed legislation in 1980 that increased the frequency of health reassessments (Continuing Disability Reviews) for SSI and especially DI beneficiaries. In the subsequent three years, more than 380,000 DI beneficiaries were terminated from the program – more than 40 percent of those who had their cases reviewed (Rupp and Scott, 1998) – for no

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labor market impacts of disability programs. Our study is unique in analyzing the interaction of program liberalization, rising replacement rates, and adverse labor demand conditions in hastening the labor force exit of less-skilled workers, thereby lowering aggregate unemployment.

[2] Approximately one-fourth of DI recipients also receive funds from SSI.

longer meeting medical standards. Additionally, Congress required the Social Security Administration to tighten the medical eligibility criteria, leading to a further decline in award rates for both programs. These dramatic reductions in benefits, occurring during the deepest U.S. recession in the post-war period, were met by intense public criticism. Citing violations of due process, seventeen states officially refused to comply with the disability review process during 1983 and 1984.

In response, Congress passed legislation in 1984 that profoundly altered the application process for both DI and SSI, yielding a more expansive definition of disability and permitting a greater exercise of voice by applicants and their medical providers.[3] SSI awards almost doubled from 1982 to 1986 while the number of DI awards increased by 38 percent during the same time period. The Continuing Disability Review process came to an almost complete halt; in the five years from 1985 through 1989, fewer individuals were terminated for no longer meeting medical eligibility standards than were terminated in the first five months of 1982.[4]

Figure 1 shows the decline in disability beneficiaries accompanying the early 1980s clampdown and the substantial increase following the 1984 liberalizations. Between 1984 and 1999, the number of adults receiving disability benefits increased at an average annual rate of 4.6 percent, with 1999 cash benefits paid out exceeding $70 billion.[5] Also noteworthy in the figure is the fluctuating mortality rate of disability recipients, which appears to inversely track program expansions.

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[3] SSA was required to: 1) relax its strict screening of mental illness by placing less weight on diagnostic and medical factors and relatively more on functional factors; 2) give highest weight to the source evidence provided by the applicant's own health care provider above the SSA's own consultative exam; 3) give more consideration to pain and related factors; 4) consider multiple non-severe impairments as constituting a disability (whereas prior to 1984 applicant were required to have at least one severe impairment; 5) desist in terminating benefits for any individual for whom SSA could not demonstrate compelling evidence of health improvement; 6) provide benefits for those former recipients whose terminations were under appeal; and 7) suspend Continuing Disability Reviews for mental impairments and pain until appropriate guidelines could be developed. See Stapleton, et al (1998) for a detailed discussion.

[4] In the post-1984 period, Congress made two additional changes that expanded the number of beneficiaries: in 1989, Congress directed SSA to conduct outreach to potentially eligible low-income individuals to inform them about SSI benefits; and in 1991, Congress required SSA to place even more weight on the information provided by an SSI or DI applicant's own medical provider. An exception to this expansionary trend was Congress' 1996 discontinuation of benefits for individuals who qualified for disability on the basis of alcohol and drug addiction, resulting in the termination of approximately 130,000 beneficiaries (cf., Lewin, 1998).

b. Changes in the characteristics of DI and SSI recipients

Accompanying the growth of the disability beneficiary population after 1984 are four demographic shifts. First, as is visible in Figure 2, DI award rates for those under the age of forty more than doubled from 1979 to 1993, while the corresponding rate for those aged 55-64 remained virtually unchanged. Accordingly, the age of new recipients declined and the fraction of new awards accounted for by those aged 40-54 rose significantly.[6]

Second, as documented in Table 1A, new beneficiaries were increasingly likely to suffer from impairments with comparatively low mortality. The share of DI awardees with a primary diagnosis of mental disorder or disease of the musculo-skeletal system – the two disorders with the lowest mortality among SSA's fourteen major diagnostic categories (Hennessey and Dykacz, 1993) – grew by sixty percent between 1983 and 1998. The corresponding shares for neoplasms and circulatory system diseases, both of which have above-average mortality, declined by 40 percent. The overall mortality rate for DI beneficiaries fell by approximately one-third from 1984 to 1999.[7]

Third, while disability recipiency rates rose for all education groups, the increase was most pronounced among high school dropouts. Data from the Survey of Income and Program Participation (SIPP) compiled in Table 1B indicate that the share of high school dropouts receiving disability benefits more than doubled between 1984 and 1999. At present, a 40 – 54 year old high school dropout is more than four times as likely to receive federal disability benefits as an individual of the same age bracket who has completed high school. As is shown in Table 1C, despite a decline in male high school dropout labor force participation of 4 percent between 1984 and 1999, the share of high school dropout *non-participants* receiving disability rose sharply, most markedly for males ages 25 – 39.

Finally, the gender composition of the DI population became substantially more representative of the

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[5] The recipiency numbers exclude both dependents of DI recipients and beneficiaries of state-only SSI benefits.

[6] Here and throughout the paper, details of our data sources and methods are provided in the Data Appendix.

[7] These changes in the age and impairment distribution of disability recipients have also substantially increased the expected benefit recipiency duration of newer cohorts (Rupp and Scott, 1998).

overall population during this interval. In 1984, there were 2.06 male DI beneficiaries per female beneficiary. By 1999, that ratio had fallen to 1.35.

c. Rising replacement rates

An overlooked factor that has likely contributed to the growth of the disability population since the 1984 liberalization is the rise in the earnings replacement rate. This rate rose steadily throughout the 1980s and 1990s due to an unforeseen but nonetheless quantitatively important interaction between the disability benefits formula and the growth of earnings inequality in the U.S. (cf., Katz and Autor, 1999).

An individual's DI benefit is a concave function of prior earnings. To determine it, SSA first calculates an individual's Average Indexed Monthly Earnings (AIME),

$$(1) \qquad AIME = \frac{1}{T}\sum_{t=1}^{T}(earn_t)\frac{avg.wage_T}{avg.wage_t},$$

roughly equal to an individual's average monthly earnings (conditional on employment) indexed by average wage growth in the U.S. economy. Benefits awarded (the Primary Insurance Amount or PIA) depend on the AIME as follows:

$$(2) \qquad PIA = \begin{cases} 0.9*AIME & if \quad AIME \in [0,b1] \\ 0.9*b1+0.32*(AIME-b1) & if \quad AIME \in (b1,b2], \\ 0.9*b1+0.32*(b2-b1)+0.15*(AIME-b2) & if \quad AIME > b2 \end{cases}$$

where the 'bend points' $(b1,b2)$ are rescaled each year by average wage growth in the economy.

As is clear from the concavity in this piecewise linear formula, low income workers replace a greater share of their earnings with disability income. More subtly, the indexation of benefits to the mean wage in the U.S. economy ensures that increases in U.S. earnings inequality benefit workers at the low end of the earnings distribution. Because the bend points rise each year with the average nominal wage, workers who experience wage growth below this average have a larger fraction of their wage income replaced on the steeper sections of the PIA formula. In addition, because each worker's potential benefits depend upon his or her entire earnings history, workers whose wages grow more slowly than the national average will have an AIME in excess of their current earnings.[8]

The distributional impacts of these attributes of the benefits formula are seen in Table 2. In 1979, male workers ages 50 – 54 at the 10th percentile of the (age-specific) earnings distribution were able to replace 53 percent of their earnings with disability income. By 1998, this number had increased to 63 percent. As the left-hand columns of Panel A demonstrate, approximately half of this increase is due to the rise in the 'bend points' relative to the lower tail of the earnings distribution, with the other half due to sluggish wage growth at the lower tail (which raises the AIME relative to current earnings). Moreover, these figures understate the growth in the total replacement rate for low wage workers since they do not account for the rising value of Medicare benefits provided to all DI recipients. By contrast, workers at the 90th percentile of the earnings distribution saw no increase in their replacement rate (excluding the value of Medicare) over this time.[9]

To summarize, the disability beneficiary population of 1999 is larger, younger, more female, less likely to suffer from high mortality disorders, and inclusive of a far larger share of the less-skilled population than the disability population of 1984. Since there is little evidence to suggest that the prevalence of disabling illness has grown substantially over this time, it is likely that these changes are in part explained by a more expansive definition of disability, changes in the determination process, and a rising replacement rate.[10]

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[8] Note that these effects are additive rather than offsetting. The earnings history effect raises the AIME, and the indexation of the bend points raises the share of the AIME replaced at a more generous rate.

[9] Details of the construction of replacement rates are given in the Data Appendix. The benefits provided to SSI recipients, though not dependent on prior earnings, have also grown rapidly over 1979 – 1998. For example, nominal SSI benefits, which are indexed to the Consumer Price Index, increased by 133 percent between 1979 – 1998, relative to wage growth at the 10th percentile of earnings of only 75 percent. We focus on DI replacement rates because the DI program is by design likely to be far more relevant to labor force participants.

[10] To be clear, this set of facts does not imply that current disability beneficiaries are undeserving. As Bound and Waidmann stress (1992), the notion that disability is a dichotomous medical state is a false one. While a more expansive definition of disability will accommodate a greater range of illness, this is a matter of societal choice rather than medical certainty. Lakdawalla, Bhattacharya and Goldman (2001) present some evidence from health self-reports that the incidence of asthma among the young has risen since 1984. However, they also note that growth in self-reported disability is entirely accounted for by less educated individuals who are leaving the labor force at greater rates; among the employed, there is no difference across education groups in the growth rate of disability. Hence, we view their findings as consistent with a story that emphasizes changes in the supply of benefits rather

2. Model

a. Structure

To guide our empirical exploration of the labor market impacts of these programmatic changes, we consider the steady state of a simple dynamic programming model of an individual's decision to apply for disability benefits. We assume that an individual has a per-period utility of work equal to $u(w)-e(h)$ where the utility of work, assumed separable in health and earnings, is increasing in the wage rate, w, and the disutility of effort, $e(h)$, is declining in health, h.[11] We initially restrict attention to the case in which neither w nor h is time varying and normalize the per-period utility of unemployment to zero.

At the beginning of each period, both employed and unemployed individuals must decide whether to apply for disability benefits. A disability applicant with health h will qualify for benefits with probability $p = p(h)$, and we assume that this person can reapply for benefits if his application is rejected. An unemployed worker who searches for a job will find one with probability q. Hence the value function for an unemployed person is:

(3) $$V_U = \max[\beta p V_D + \beta(1-p)V_U,\ q\beta V_E + \beta(1-q)V_U],$$

where β is the discount rate and V_E and V_D represent the value of employment and disability recipiency, respectively.[12] The first term is the expected utility of applying for disability benefits while the second equals the corresponding value associated with searching for a job.

Because the parameters in this equation are not time-varying, an individual's optimal decision rule is unchanged in the next period if he remains unemployed and thus the preceding equation simplifies to:

(4) $$V_U = \max\left[\frac{\beta P V_D}{1-\beta(1-p)}, \frac{\beta q V_E}{1-\beta(1-q)}\right]$$

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than shifts in underlying health. Bound and Waidmann (2000) also stress that the incidence of self-reported disability among males responds markedly to changes in the generosity of the disability program.

[11] In modeling the disability application decision as a function of both health and the disutility of work, we follow the approach of Diamond and Sheshinski (1995). See also Hausman and Halpern (1986) and Benitez-Silva et al. (2000) for theoretical and empirical analyses of the application decision.

All else equal, an unemployed individual will be more likely to apply for disability benefits if her probability of qualifying is higher, the program is more generous, or the reemployment odds are lower.

Employed workers will choose between working and applying for disability benefits. Consistent with the eligibility requirements, we assume that an individual cannot work while applying for the program and therefore must exit employment (either voluntarily or involuntarily) to apply. If an individual chooses to remain employed, she receives a net utility of $u(w)-e(h)$ in the current period and retains her job in the subsequent period with probability $(1-s)$. If instead she finds it optimal to exit employment to apply for disability benefits, then she will also find it optimal to re-apply while unemployed if her application is rejected. Once the application is accepted, she will not voluntarily leave the program. Hence, the asset value of disability recipiency, V_D, simplifies to $V_D = u(d)/(1-\beta)$, with $u(d)$ representing the per-period utility of receiving disability benefits, d. The value of applying for disability is therefore equal to V_D discounted by the expected time to an award. The value of employment is thus:

$$(5) \qquad V_E = \max\left[u(w)-e(h)+\beta\big((1-s)V_E+sV_U\big),\ \frac{\beta P V_D}{1-\beta(1-p)} \right]$$

Notice that the value of the first term will depend on whether in the event of job loss, an individual chooses to search for work or to apply for disability benefits.

b. Comparison of steady states

Using this structure, we compare the steady states of the model under different labor market (q and s) and program generosity conditions (d and p). We are particularly interested in two questions: How do these parameters impact the propensity of workers to leave employment to apply for disability benefits? And how do they impact the probability that *conditional* on job loss, workers leave the unemployment pool to seek benefits? We refer to the first impact as the 'level effect' of the disability program on labor force participation. We refer to the second as the 'interaction' between the disability

[12] Thc constant discount rate implicitly captures both time discounting and a constant mortality risk.

program and labor demand conditions, since this force is only operative when workers experience adverse employment shocks.

Given the structure of the disability application process, employed workers will optimally choose one of three actions in each period: (1) quit employment to apply for disability; (2) remain in employment but apply for disability *conditional* on unemployment and (3) remain in the labor force and not seek disability benefits. We label these groups (a), (c), and (n) respectively (for always apply, conditionally apply, and never apply). Using equation (4) and the first term in equation (5) to solve for the value of employment for groups (c) and (n), the resulting expression for $V_E = \max[V_{ea}, V_{ec}, V_{en}]$ is:

$$(6) \qquad V_E = \max\left[\frac{\beta p V_D}{1-\beta(1-p)}, \frac{v(1-\beta(1-p))+\beta p s V_D}{(1-\beta(1-s))(1-\beta(1-p))}, \frac{v(1-\beta(1-p))}{(1-\beta)(1-\beta(1-q-s))}\right]$$

Individuals with a low net utility of work – either because of low wages, ill health, or some combination of the two – are likely to fall into category (a), while the (n) group will be made up of relatively high wage earners and/or those in good health. Finally, group (c) consists of "conditional applicants" – individuals who choose to remain in their job but who optimally apply for disability benefits rather than search for a job once they become unemployed.

Figure 3 displays a simulation of the value of these three actions as a function of an individual's per period net utility of work.[13] As is clear from the figure, the value of employment to the non-applicants (n) rises more steeply with increases in $u(w)-e(h)$ than does the corresponding value for conditional applicants (c). Logically, the value of exiting the labor force to apply for benefits is independent of $u(w)-e(h)$; once a worker exits the labor force, she will optimally choose not to return. The key point underscored by this figure is that that the value of employment for individuals at differing levels of $u(w)-e(h)$, is equal to the upper envelope of these three value functions (denoted by the dashed line in Figure 3). Consequently, the salient question for our analysis is how do the parameters of the disability

[13] Parameter values for the simulation are $\beta = 0.9$, $p = 0.5$, $u(d) = 1$, $s = 0.1$, and $q = 0.5$.

program and the state of the labor market impact the proportion of workers pursuing each action.

Define $\tilde{u}_{AC}$ as the net utility of work $u(w)-e(h)$ at which an individual is indifferent between exiting the labor force immediately to apply for benefits and applying for benefits only in the event of job loss: $V_{ea}(\tilde{u}_{AC}) = V_{ec}(\tilde{u}_{AC})$. Similarly, define $\tilde{u}_{CN}$ as the net utility of work at which an individual is indifferent between applying in the event of job loss and never applying.[14] Solving for $\tilde{u}_{AC}$ and $\tilde{u}_{CN}$ yields:

$$(7) \qquad \tilde{u}_{AC} = \left(\frac{\beta p}{1-\beta(1-p)}\right)u(d), \qquad \tilde{u}_{CN} = \left(\frac{pu(d)}{q}\right)\left(\frac{1-\beta(1-q-s)}{1-\beta(1-p)}\right).$$

Factors that move $\tilde{u}_{CN}$ rightward will increase the size of the "conditional applicant" group at the expense of the "never apply" group. Factors that move $\tilde{u}_{AC}$ rightward will increase the size of the "always apply" group at the expense of the "conditional applicant" group. Factors that move both $\tilde{u}_{CN}$ and $\tilde{u}_{AC}$ rightward will increase the size of the total applicant pool and shrink the size of the non-applicant pool, while leaving the net impact on the size of the "conditional group" ambiguous.

Initially, we consider how labor market conditions (q and s) affect equilibrium behavior. Observe that $\tilde{u}_{AC}$ does not depend on either the job loss or reemployment odds. To see the intuition for this result, consider an individual who is indifferent between immediate and delayed (until job loss) disability application. Indifference at $\tilde{u}_{AC}$ implies that the per-period utility of employment is identical to the expected per-period utility of disability application. Accordingly, changes in the rate of job loss, s, that hasten or delay the moment of disability application have no impact on $\tilde{u}_{AC}$. Furthermore, as noted above, once an individual leaves the labor market to apply for benefits, it is never optimal for her to return, implying that $\tilde{u}_{AC}$ is also independent of q.

Hence, comparing across steady states of the model, it is apparent that changes in labor market

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[14] Note we are implicitly assuming that p is constant. In reality, p will depend upon h. It is therefore useful to consider this exercise as applying to a given individual with health h and $p(h) > 0$. In this case, Figure 3 depicts an individual's optimal decision rule for differing values of w, holding h_i constant.

conditions do not affect the size of the group that directly exits current employment to seek benefits.[15] Instead, adverse shifts raise the option value of disability application, thereby increasing the size of the conditional applicant group. Because increases in s and declines in q reduce the value of job search, it is straightforward to show that the search/apply utility threshold, $\tilde{u}_{CN}$, increases with s and falls with q. Pictorially, higher s and/or lower q move $\tilde{u}_{CN}$ in Figure 3 rightward while leaving $\tilde{u}_{AC}$ unaffected. Consequently, poorer labor market conditions among the potential applicant population (i.e., $p(h)>0$) unambiguously increase the share of workers who choose to exercise the option of applying for benefits in the event of job loss.[16]

Now consider an increase in the generosity of the disability program, either through benefit increases or reductions in screening stringency – both of which are likely to have occurred between 1984 and 1999. Logically, increases in d and p shift both $\tilde{u}_{AC}$ and $\tilde{u}_{CN}$ rightward, increasing the fraction of individuals who apply for benefits. As the disability program becomes more generous, the share of employed and unemployed seeking benefits increases unambiguously and the corresponding share that never applies declines. Hence, increases in program generosity exert a 'level effect' on labor force participation by directly inducing some workers to leave employment to seek benefits. Moreover, it is straightforward to show that $\partial\tilde{u}_{CN}/\partial d > \partial\tilde{u}_{AC}/\partial d$ and $\partial\tilde{u}_{CN}/\partial p > \partial\tilde{u}_{AC}/\partial p$. That is, increases in d and p shift the conditional/never margin, $\tilde{u}_{CN}$, farther rightward than they shift the always/conditional margin, $\tilde{u}_{AC}$. This

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[15] Of course, higher s implies greater flows into the disability applicant pool. Note that the result that $\tilde{u}_{AC}$ is independent of s and q derives in part from our assumption that w is fixed – workers keeps their wages until job loss. More realistically, if w varies with s and q, deterioration in labor market conditions may induce some individuals to exit employment to apply for benefits. While our assumption that w is constant is clearly too stark, it is qualitatively consistent with the empirical evidence that wages of incumbent workers are substantially more sheltered from labor market conditions than are new market entrants (cf., Beaudry and DiNardo, 1991).

[16] The responsiveness of the size of the group in the region $\tilde{u}_{AC} - \tilde{u}_{CN}$ to changes in either q or s is likely to be significant. For the parameter values used in Figure 3, an increase in s from 0.10 to 0.15 increases the width of this region by more than 23 percent, while a reduction in q from 0.50 to 0.40 increases it by approximately 25 percent. If, for example, the distribution of net utilities was uniform, these changes would induce correspondingly large increases in conditional applicants flows.

suggests that the size of the conditional applicant group will also grow in response to an increase in program generosity, although this prediction is formally ambiguous without further assumptions on the distribution of net utilities.[17]

Finally, we find that the level of generosity of the disability program interacts positively with adverse labor market conditions to increase the size of the conditional applicant group. More formally, the cross-partial derivatives $\partial^2\tilde{u}_{NC}/\partial d\partial s$ and $\partial^2\tilde{u}_{NC}/\partial p\partial s$ are strictly positive while $\partial^2\tilde{u}_{NC}/\partial d\partial q$ and $\partial^2\tilde{u}_{NC}/\partial p\partial q$ are strictly negative. By contrast, all four corresponding cross-partial derivatives for the always/conditional threshold, $\tilde{u}_{AC}$, are zero. Hence, the more generous are program benefits or the less stringent is program screening, the more that adverse labor market conditions increase the size of the conditional group.[18]

c. Time varying wages and health

While our framework makes the implausible assumption that neither wages nor health are stochastic, there are three reasons to think that dynamic considerations reinforce our conclusion that increasing program generosity raises the propensity of the unemployed to exit the labor force. First, it is likely that the expectation of w for an individual declines with job loss (Jacobson, LaLonde and Sullivan, 1993), thereby raising the attractiveness of disability to the unemployed relative to the employed. Second, the calculation of benefits under the disability program ensures that a worker whose wages are growing more slowly than the national average who remains in his job will enjoy a substantial gain in annual disability payments because of the indexation of past wages. Finally, to the extent that health deteriorates over time for many individuals who consider applying for this program, there is a gain associated with waiting

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[17] This implication will hold provided that (1) the density function of net utilities is weakly monotonically increasing in $u(w)-e(h)$ below its mean and (2) conditional on p and d, $\tilde{u}_{AC}$ and $\tilde{u}_{CN}$ are both below the mean of the net utility distribution. Many distributions, including the normal and uniform, satisfy (1).

[18] Of course, after sufficient time had elapsed in the model, all "conditional applicants" would have exited the labor force, at which point the program would exert no further effect on their unemployment propensity. Hence, it is useful to think of the model as applying to a single cohort of workers, with new cohorts entering the market continuously. Note that individuals must typically participate in the labor market for several years to qualify for DI benefits.

because of an increase in the probability of qualifying for benefits.

Hence, our model provides ample reason to suspect that liberalization of the disability program – that is, lower screening stringency and rising replacement rates – coupled with declining labor market prospects for the low skilled (Juhn, 1992; Juhn, Murhpy and Topel, 1991) are likely to have both decreased labor force participation and increased the propensity of job losers to exit the labor force to seek disability benefits.

3. Disability and labor force participation: Level effects

We begin the empirical analysis by exploiting the disability retrenchment of 1979 – 1984 and subsequent liberalization over 1984 – 1998 as a quasi-experiment for studying the impact of disability benefits on the *level* of labor force participation. The equation we would initially like to estimate is:

(8) $$P[LFP_i] = P[\alpha + \beta_1 g(REP_i, p(h_i)) + \beta_2 W_i + \beta_3 h_i + X_i \beta_4 > -\varepsilon_i],$$

where LFP_i is a dichotomous variable equal to one if individual (i) is a labor force participant, W_i is the market wage, h_i is individual health, X_i is a vector of individual characteristics, ε is a normally distributed error term, and $g(REP_i, p(h_i))$ is a DI benefits 'supply function' which is increasing in an individual's replacement rate, REP_i, and in her probability of obtaining benefits conditional on application, $p(h_i)$. In this equation, we expect that $\beta_1 < 0$ and $\beta_2, \beta_3 > 0$.

There are a number of problems in estimating this equation, however. First, we cannot simultaneously observe REP_i and W_i for a given individual. Second, objective measures of h_i (and $p(h_i)$) are typically unavailable from survey data sources. And, finally, as stressed by Bound (1989), because individuals with poor health are likely to command lower wages, omission of h_i from (8) will bias estimates of β_2 and to the degree that REP_i and h_i are correlated with W_i, β_1 as well.

To surmount these obstacles, we estimate a state level analog of (8) in first differences,

(9) $$\Delta LFP_{jT} = \alpha + \beta_1 \Delta g(\cdot)_{jT} + \beta_2 \Delta W_{jT} + \beta_3 \Delta h_{jT} + \Delta X_{jT} \beta_4 + \varepsilon_{jT},$$

where (j) subscripts the 50 U.S. states excluding the District of Columbia and Δ denotes the first difference operator. As an empirical analog to $\Delta g(\cdot)$, we initially use the observed contemporaneous state level changes in DI and SSI recipiency (ΔDI_{jT}, ΔSSI_{jT}). Subsequently, we apply an instrumentation strategy to circumvent possible endogeneity in this measure. Estimates of (9) also control for changes in the age structure by state-education group. Since conditional on age and education average wage and health changes are likely to be common across states, we do not control for these variables directly but rather allow them to be absorbed by α .

a. OLS Estimates

Table 3 presents initial estimates of (8). We perform separate estimates for high school dropouts and those who have completed high school. The first specification, which parameterizes disability recipiency as the sum of the change in DI and SSI recipiency, indicates that changes in disability recipiency from 1979 - 1984 are significantly negatively related to contemporaneous changes in the labor force participation of high school dropout males. A one percentage point increase in disability recipiency is predicted to reduce male high school dropout labor force participation by 3.2 percentage points. Since high school dropouts compose a far larger share of the disability recipient population than their share in the workforce, this coefficient is of reasonable magnitude.[19] Notably, when changes in DI and SSI recipiency are entered separately into this equation, it is only the change in DI recipiency that is significantly related to the change in labor force participation.[20]

Columns (3) and (4) display analogous estimates for the labor force participation of high school dropout males during 1984 – 1998. Despite the fact that the disability program was contracting rapidly during 1979 – 1984 and expanding thereafter, the estimated impacts of disability recipiency levels on labor force participation are quite similar in the post-reform era. Hence, male high school dropouts were differentially *entering* the labor force in states with large reductions in disability recipiency during the

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[19] Using CPS data, we calculate that male high school dropouts composed on average 7.9 percent of the labor force during 1984 to 1998. SIPP data indicate that the male high school dropout share of all disability recipients was approximately 25 percent in 1999.

[20] This finding is logical since DI recipients must have several years of work experience in covered employment to

retrenchment and differentially *exiting* in high disability growth states following the liberalization.

Columns (5) – (8) of Table 4 display comparable specifications for the labor force participation of males with a high school education or greater. Consistent with the patterns in Table 1, the estimated relationships between changes in disability recipiency and changes in male labor force participation are much weaker for more educated workers and are generally insignificant.

As noted above, the absolute growth in DI recipiency per population was slightly larger among females than males over 1984 – 1998. Hence, we provide estimates in Panel B of Table 3 comparable to those above for the labor force participation of women. Although the point estimates are less precise here than in the corresponding male estimates, we again find that DI but not SSI recipiency is negatively related to the labor force participation of high school dropouts during both the retrenchment and reform periods. No similar relationship is detected for females with a high school education or greater.

b. Instrumental Variables Estimates

A concern with the Table 3 estimates is that growth in disability benefits at the state level may be driven by changes in the 'supply' of benefits – due to the program retrenchment and reform – and by changes in the 'demand' for benefits due to earnings and health shocks. Depending on the cross-state correlation between supply and demand shocks, the Table 3 estimates may be biased either upward or downward. To address this concern, we note that the key intervening variable that potentially links labor force participation and disability recipiency is disability application: an individual cannot receive disability without applying for benefits and cannot apply for benefits while employed. Hence, factors that exogenously impact the supply of disability benefits will induce corresponding changes in the flow of applications, which can then be used to identify the characteristics of the marginal group of applicants – in particular the extent to which they are drawn from the labor force.

Concretely, consider a variant of (8) where the dichotomous outcome variable is an individual's decision to apply for disability benefits:

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qualify for DI benefits. No work history is required to qualify for SSI.

(10) $P[Apply_i] = P[\alpha + \beta_1 REP_i + \beta_2 P(h_i) + \beta_3 REP_i \cdot P(h_i) + \beta_3 W_i + \beta_4 h_i + X_i \beta_5 > -\mu_i]$.

In this equation, we have taken a linear approximation to the 'disability supply function' $g(REP_i, p(h_i))$ where $\hat{\beta}_1 = g_1(\cdot)$, $\hat{\beta}_2 = g_2(\cdot)$ and $\hat{\beta}_3 = g_{12}(\cdot)$. Our expectation is that $\beta_1, \beta_2, \beta_3 > 0$.

Aggregating (10) to the state level and taking first differences, we obtain:

$$\begin{aligned} \Delta Apply_{jT} &= \alpha + \beta_1 \Delta REP_{jT} + \beta_2 \Delta P(h_{jT}) + \beta_3 [REP_{jT} P_T(h_{jT}) - REP_{jt} P_t(h_{jt})] \\ &\quad + \beta_4 \Delta W_{jT} + \beta_5 \Delta h_{jT} + \Delta X_{jT} \beta_6 + \varepsilon_{jT} \end{aligned} \tag{11}$$

where $\Delta Apply_{jT}$ is the change in the application flow over two points in time as a function of the supply and demand for benefits. Note that in this equation, we add a time subscript to the screening function $P_t(h)$ to indicate that screening stringency differs between the retrenchment and reform periods.

To obtain exogenous variation in the *supply* of benefits, we exploit the fact that the DI benefits formula is progressive but is not indexed to regional wage levels. As a result, workers in low wage states face significantly higher earnings replacement rates.[21] Consistent with this fact, workers in high replacement rate states have higher rates of disability application and recipiency at a point in time, as is visible in Panel A of Figure 4. This state level variation in DI replacement rates provides us with two candidate instruments for identifying plausibly exogenous cross-state shifts in the supply of DI benefits. First, during the retrenchment of 1979 – 1984, SSA applicant screening stringency rose substantially (i.e., $\Delta P_T(h_j) < 0$). Provided that $\beta_3 \neq 0$ – that is, the impact of screening stringency on disability applications varies with the replacement rate – the *interaction* of increasing screening stringency with cross-state variation in replacement rates will induce differential state level reductions in the supply of DI benefits over 1979 – 1984.[22] This relationship is readily detectable in Panels B and C of Figure 4, which plot simulated state 1979 replacement rates against ensuing state disability population reductions and

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[21] A regression of mean log male high school dropouts wages by state in 1984 on the simulated replacement rate variable and a constant yields a coefficient of –4.31 with t-ratio of 9.0 ($R^2 = 0.63$).

[22] Note that we assume that $\Delta P_t(h_j)$ is approximately uniform across states and hence will be absorbed by α in (11).

expansions during 1978 – 84 and 1984 – 98.

Second, as discussed above, rising U.S. earnings dispersion induced substantial increases in effective DI replacement rates over 1979 – 1998. This effect was not uniform across states, however. Because workers in low wage states replace a larger fraction of their earnings on the steeper sections of the progressive earnings replacement formula, effective replacement rates rose by considerably more in low wage states. This exogenous variation in ΔREP_{jT} will again induce differential state level increases in the supply of DI benefits over 1984 – 1998 provided that $\beta_1 > 0$.[23]

To implement this instrumental variables approach empirically, we estimate a version of (9) where we replace $\Delta g(\cdot)_{jT}$ with $\Delta Apply_{jT}$, the change in the state level flow of disability applications, and instrument this variable using levels and changes in projected state level DI replacement rates. To form an estimate of the change in the state level flow of applications, $\Delta Apply_{jT}$, we multiply the difference in the level of DI applications per populations at the start and end years of a time interval (i.e., 1979 – 84, 1984 – 98) by the number of elapsed years. We estimate state level replacement rates and their changes as a function both of the increasing effective generosity of DI benefits (equations (1) and (2)) and the changing age distribution in each state. As above, we control for state level changes in the age distribution of the relevant gender-education group.

An additional complication for our estimation is that over 1979 – 1984, a large number of disability beneficiaries – approximately 560 thousand – were terminated during the Continuing Disability Review process for no longer meeting medical eligibility criteria. Although the fraction of *recipients* terminated was likely to be roughly equivalent across states, these reductions may imply proportionately larger reductions in the fraction of state *residents* receiving benefits in high replacement rate states (where the initial recipiency rate was far higher). On the assumption that a share of terminated beneficiaries is likely

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[23] In theory, it would also be possible to use cross-state changes in ΔREP_{jT} over 1979 – 1984 to instrument changes in the supply of benefits. In practice, we are not able to estimate these cross-state shifts with sufficient precision over this short term interval.

to have reentered the labor force in these states (cf., Bound, 1989), failing to account for the terminated beneficiaries would cause us to overestimate the labor force participation propensity of the marginal applicant.[24] To handle this potential bias, we make the conservative assumption that the marginal terminated beneficiary has one half the labor force participation propensity of the marginal applicant.[25] Accordingly, we add to $\Delta Apply_{j,79-84}$ one half of the number of estimated state level beneficiaries terminated per state population during these years.[26]

Table 4 presents OLS and instrumental variables estimates of the relationship between DI applications and labor force participation of high school dropouts and high school completers by gender for 1979 – 84 and 1984 – 1998. Column (1) of Panel A presents an initial OLS estimate for male high school dropouts. Over 1979 – 1984, reductions in DI applications are associated with increases in labor force participation of male high school dropouts: a 1 percentage point increase in the rate of disability applications per state population is associated with a 1.66 percentage point decline in male high school dropout labor force participation.

Because of the concern that this estimate may be biased by the endogeneity of the applications variable, we instrument the change in applications, $\Delta Apply_{j,79-84}$, by the state level replacement rate in 1979.[27] The coefficient of –5.23 on the instrumented application rate variable indicates that a one-half

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[24] This follows because the reduction in applicants per population and the share of beneficiaries terminated per population over 1979 – 1984 were both positively correlated with state level replacement rates. Our data indicate that applications fell by approximately 7 per thousand state residents over 1979 – 84. During the same period, the number of beneficiaries terminated was approximately 5 per thousand state residents.

[25] Bound (1989) estimates that approximately 50 percent of rejected disability applicants over 1972 – 1998 re-entered the labor force. Assuming this estimate applies to the marginal applicants in our sample, our assumption implies that 25 percent of terminated beneficiaries would re-enter the labor force.

[26] It is important to note that the number of beneficiaries terminated by state is not available from SSA. To impute this number, we allocate the national count of terminations over 1979 – 1984 to states according to their observed reduction in DI applications per population over 1979 – 1984. This adjustment does not qualitatively change the sign or statistical significance of our findings, but does of course reduce (by about 30 percent) the implied impact of exogenous reductions in DI applications on labor force participation. A variety of alternative allocation schemes – for example, using changes in recipient count or initial recipiency as allocation weights – lead to near identical results.

[27] The first stage coefficient is given in the footer of each column of the table. In all cases, the first stage is highly significant. The magnitude of the first stage coefficient (-0.025) does not have a clear structural interpretation since

percentage point decrease in applications per population induced by the growing screening stringency during 1979 – 1984 caused a corresponding increase in high school dropout male labor force participation of approximately 2.5 percentage points. Note that this impact is of reasonable magnitude: over 1979 – 1984, approximately one-eight of all labor force participants were male high school dropouts, relative to fifty percent of disability recipients.

Columns (3) – (5) of Table 4 present similar estimates for 1984 – 1998. These models confirm that an exogenous increase in the supply of DI benefits increases the flow of DI applications and reduces the labor force participation of male high school dropouts. The point estimates for these models are comparable to albeit slightly smaller in magnitude than the instrumental variables estimates for 1979 – 84. One noteworthy feature of these estimates is that our instruments in the pre-1984 versus post-1984 period use different sources of variation (levels versus changes of the replacement rate) and have opposite impacts on the level of DI applications. Yet the implied impact of increases in the supply of benefits on the labor force participation of high school dropout males is comparable across periods. This enhances our confidence that the findings are not artifactual or idiosyncratic to a single episode of disability program reform.

A reduced form version of these relationships is depicted in Figure 5. This figure plots 1979 state replacement rates against levels and changes of male high school dropout labor force participation for 1979 to 1998. Consistent with the IV estimates, the figure underscores that the interaction of cross-state variation in replacement rates with changes in program stringency exerted a noticeable effect on participation rates of high school dropouts over both 1979 – 84 and 1984 – 98.

Panel B of Table 4 presents analogous estimates for female high school dropouts. Although the instrumental variables coefficients are only marginally significant for females, the point estimates again suggest that exogenous increases in the supply of disability benefits reduce the labor force participation of

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it measures the impact on applications of the *interaction* between the state replacement rate and an unobserved variable, the national change in screening stringency for marginal applicants.

less skilled females.[28] Consistent with earlier findings, the OLS and IV estimates detect no significant impact of the changing supply of disability benefits on the labor force participation of males and females with a high school degree or greater.

To summarize, the data suggest that the stringency and generosity of the disability regime has an important direct effect on the labor force participation of low-skilled individuals, particularly high school dropouts of both genders. The point estimates indicate that the disability retrenchment during 1979 – 84 raised male high school dropout labor force participation by as much as 4.7 percentage points in high replacement relative to low replacement rate states. And during the ensuing 14 years, liberalization of the disability program induced a similarly large relative percentage point drop in high school dropout labor force participation among high replacement states – on the order of 7.2 percentage points.[29] Though less precisely estimated, the point estimates suggest that the labor force impacts on female high school dropouts were about two-thirds as large.

4. The interaction of disability with adverse employment shocks

Because the unemployed intrinsically face lower opportunity costs of labor force withdrawal, the disability program is likely to differentially attract applications from the unemployed. Moreover, as our model indicates, reductions in disability screening stringency, increases in program generosity, and deterioration in labor market conditions all serve to increase the relative attractiveness of disability application for the unemployed relative to the employed – raising the share likely to exit the labor force in

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[28] It is not altogether surprising that the simulated replacement rate instrument has essentially no explanatory power for the labor force participation rate of female high school dropouts since it is calculated for males only. Given the rapidly shifting patterns of labor force participation among women during this period, calculating comparable earnings simulations for women presents a significant challenge that we have not undertaken.

[29] For example, the simulated replacement rate was 7.26 percentage points higher in Arkansas than Washington in 1979, which implies a differential decline in applications per population of 0.90 percentage points (100*(.0726)*(-0.025)*5) over 1979 – 1984. Using the 2SLS coefficient from column (2) of Table 4, this implies a relative increase in male high school dropout labor force participation of 4.7 percentage points (0.90*5.23) in Arkansas versus Washington during these five years. Between 1984 and 1998, the simulated replacement rate rose by 1.1 percentage points more in Arkansas than Washington (4.92 versus 3.82 percentage points), implying a differential increase in DI applications per population of 2.2 percentage points (100*(0.011*2.005) over 1984 – 1998, and a differential decline in male high school dropout labor force participation of 7.2 percentage points (0.022*(–3.25)).

the event of job loss.[30]

To explore the whether the disability program has indeed served to induce labor force exit among job losers, we first ask whether the responsiveness of disability application to plausibly demand exogenous shocks has risen secularly since the disability reforms of 1984. Next, we measure whether the probability that a worker exits the labor force *conditional* on job loss has increased commensurately. We finally perform simple calculations to explore whether the increase in application propensity could plausibly explain the change in the conditional probability of labor force exit that we observe in the data. Our approach in this section is similar in spirit to Black, Daniel and Sanders (1998) who study the impact of shocks to coal prices on SSI income in mining intensive counties.[31]

a. Labor demand shocks

To implement these tests, we require a measure of plausibly exogenous labor demand shocks. Following the approach developed by Bartik (1991) and employed by Blanchard and Katz (1991) and Bound and Holzer (2000), we exploit cross-state differences in industrial composition and national-level changes in employment to predict individual state employment growth. Specifically, we calculate the predicted log employment change $\hat{\eta}_{jT}$ for each state (j) between years (t) and (T) as:

$$(12) \qquad \hat{\eta}_{jT} = \sum_k \gamma_{jkt} \cdot \eta_{-jkT}$$

where η_{-jkT} is the log change in 2-digit industry (k's) employment share nationally between (t) and (T) and γ_{jkt} is the share of state employment in industry (k) in state (j) in the initial year (t). The subscript $-j$ in η_{-jkT} indicates that each state's industry (k) employment is excluded when calculating the national employment share change.[32]

[30] Using Panel Study of Income Dynamics data for 1968 to 1991, Daly (1998) finds that 50 percent of DI recipients experienced job loss in the 5 years prior to receiving benefits. Her analysis does not explore whether the probability of applying for benefits conditional on job loss has become more prevalent with time.

[31] See also Lewin-VHI, Inc. (1995), Rupp and Stapleton (1995) and Stapleton et al. (1998) for evidence on the importance of the economic climate to disability application and recipiency.

[32] In excluding own state employment, our projected employment changes differ from those used by the authors

This methodology predicts what each state's change in employment would be if industry level employment changes occurred uniformly across states and state-level industrial composition was fixed in the short term. Accordingly, states with a relatively large share of workers in declining industries will have predicted employment declines, while those states differentially employing workers in growing industries will have predicted increases. Provided that national industry growth rates (excluding own state industry employment) are uncorrelated with state level labor supply shocks, this approach will identify plausibly exogenous variation in state employment.

b. The impact of labor demand shocks on disability applications and recipiency

As above, we use state level data to estimate an aggregate version of an individual decision problem:

$$(13) \qquad P[Apply_i \mid Job\,Loss_i] = P[\alpha + \beta_1 g(REP_i, P(h_i)) + B_3\beta_2 W_i + \beta_4 h_i + X_i\beta_5 > -\varepsilon_i],$$

where $g(\cdot)$ represents the 'disability supply' function as above. The coefficient of interest is β_1, the impact of the supply of disability benefits on the probability that an individual seeks benefits in the event of job loss.[33]

To implement this equation at the state level, we estimate:

$$(14) \qquad (APPS/POP)_{jT} = \alpha + \gamma g(\cdot)\hat{\eta}_{jT} + \beta_1 \Delta W_T + \beta_2 \Delta h_{jT} + \Delta X_{jT}\beta_3 + \delta_T + \varepsilon_{jT},$$

where the dependent variable is the cumulative count of unique DI applications per population in state (j) between the years t and T, $\hat{\eta}_{jT}$ is the contemporaneous predicted state log relative demand shock from (12), δ_T is a vector of time dummies, and other variables are defined as above.[34] In practice, we do not observe Δh_{jT}, but we assume that health shocks are unlikely to be correlated with $\hat{\eta}_{jT}$ and hence their

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cited earlier. We found that including own-state employment substantially increased the predictive power of the employment projections, a pattern that left us concerned about a potential mechanical relationship.

[33] Note that implicit in this equation is the condition that individual (i) did not *already* exit the labor force to apply for disability prior to job loss (i.e., the 'main effect' of disability on labor force participation). Hence, this equation specifically measures the interaction of disability with job loss in inducing labor force exit.

[34] Note that because $(APPS/POP)_{jT}$ measures the inflow of new applicants, it is intrinsically a flow variable – conceptually similar to specifying the change in the DI recipiency rate (ΔDI_{jT}) as the dependent variable.

omission will not bias estimates of γ.[35] Since there are conceptual problems in measuring ΔW_{jT} at the state level, we omit it from the estimates but return to the question of regional variability in wages in section 5 below. As with the prior estimates, $g(\cdot)_T$ is not observed directly, but will not bias the estimates provided that it is approximately constant across states.[36] Hence, this approach allows us to identify $\hat{\gamma} \approx g(\cdot)_T \cdot \gamma$. The key empirical prediction of our model is that because $g(\cdot)_{79-84} < g(\cdot)_{84-98}$, we have $\hat{\gamma}_{84-98} > \hat{\gamma}_{79-84}$, i.e., the demand shock-disability application locus should become steeper with time.

Before turning to estimates, Figure 6 presents a plot of the unconditional relationship between projected demand shocks and state level DI application flows during four, five-year sub-intervals of 1979 – 1998.[37] These panels reveal two key patterns. First, there is a robustly significant relationship between plausibly exogenous state level demand shocks and observed disability applications in each period. Second, the responsiveness of application rates to demand shocks rises secularly in each subsequent five-year sub-period. In the final panel, the slope has increase five-fold.[38]

To explore these relationships more formally, Table 5 presents estimates of equation (14) for 1978 – 84 and 1984 – 98. We estimate this equation using observations spaced at three-year intervals to increase precision and reduce possible serial correlation in the state level demand shock measure.[39] The first

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[35] Ruhm (2000) presents evidence that adverse demand shocks may lead to *positive* health shocks, which would therefore work against finding that adverse demand shocks increase the number of disabled.

[36] In addition, the main effect of $g(\cdot)_T$ will be absorbed by time dummies. Although we could also attempt to estimate $\hat{g}(\cdot)$ by state, the over-time variation in $\hat{g}(\cdot)$ during 1984 – 1998 is substantially more pronounced than the cross-state variation, and hence the power of this approach in the interacted specification is low. Note that because $\hat{\eta}_{jT}$ is mean zero by construction, it does not exert a main effect.

[37] Because 1998 is the most recent year of DI application data available to us, we use one year of overlapping data in forming Panels C and D of the figure.

[38] An analogous plot of the relationship between shocks and DI recipiency per state population confirms that increases in application propensity yielded conformable increase in the population of disability recipients.

[39] For 1978 – 1984 estimates, we use changes for 1978–81 and 1981–84. For 1984 – 1998 estimates, we use changes for 1984–87, 1987–90, 1990–93, 1993–96, and (imperfectly) 1995–98. All models include year dummies and controls for the education, gender, and age composition of state populations in each year. So, for example, for the period 1978 – 1981, the dependent variable is the sum of applications per population over 1979 to 1981, and $\hat{\eta}_{jt}$ is the predicted state level log employment change for 1978 – 1981. We experimented with using observations at one,

column indicates that during 1978 – 84, a one percentage point log employment shock induced a 0.086 percentage point increase in the share of all non-elderly state residents applying for DI benefits. The corresponding estimate for 1984 – 98 in column (2) reveals that a one percentage point employment shock induced more than twice as large an application response in the post-reform era.

Columns (3) and (4) repeat these estimates while pooling the data for 1978 – 1998 and adding an interaction between the shock measure, $\hat{\eta}_{jT}$, and a post-1984 dummy variable. The increase in $\hat{\gamma}$ from the retrenchment to the reform era is large and precisely estimated in column (3). The addition of state fixed effects to the model in Column (4) reduces the estimated impact of predicted demand shocks on application rates in both periods, but the increase in application responsiveness remains large and significant.[40] Panel B repeats the estimates for disability awards per population. Although awards are also significantly more responsive to demand shocks in the post-reform era, these estimates should be treated with some care. Since many applications from the mid-1990s onward were still pending at the time our data were prepared by the Social Security Administration, award rates fall off artificially in later years.[41]

We have performed numerous robustness tests to verify these relationships. If we replace the disability application variable with the change in DI recipients per population, we find qualitatively similar patterns. Further dividing the estimates into sub-intervals of 1984 – 98 indicates a secularly increasing pattern of application responsiveness, as would be expected from Figure 6. In addition, we have utilized county level DI recipiency data to explore the sensitivity of disability recipiency over 1984 – 1998 to *within-state* variation in local demand shocks generated by variation in county-level industrial structure. Even within states, we find that those counties experiencing negative predicted employment

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two, and five year intervals. The findings were qualitatively similar.

[40] To explore the possibility that given labor demand shocks lead to higher application rates in states with greater effective replacement rates, we have also performed estimates where we include the state predicted replacement rate and its interaction with the state level demand shocks. While the coefficients on the shock-generosity interaction terms are typically only marginally statistically significant, the point estimate suggests that negative employment shocks of given magnitude have a larger positive impact on disability applications where effective program generosity is higher, especially in the post-1984 period.

[41] The award rate in a given year refers to the outcome of applications *initiated* in that year, although many of the

shocks saw substantially greater growth in disability recipiency over 1984 – 1998.

c. The impact of demand shocks on labor force exit

Since effectively all disability applicants are labor force non-participants, there are two potential explanations for the findings on application responsiveness. One is that, over time, employment shocks have increasingly spurred labor force non-participants to apply for disability benefits, perhaps in anticipation of difficulty finding work in the future. A second and not mutually exclusive explanation is that the share of job losers exiting the labor force to seek disability benefits has risen in response to the 1984 reforms.

We investigate the latter possibility by estimating a state-level variant of (13) in which we use projected demand shocks to instrument for employment losses at the state level. Using these instruments, we ask whether the probability of labor force exit conditional on job loss rose between the retrenchment and reform eras. The estimating equations are:

$$(15) \qquad \Delta(NILF/POP)_{jT} = \alpha + \beta_1 \hat{\Delta}(EMP/POP)_{jT} + \Delta X_{jT}\beta_2 + \delta_T + \varepsilon_{jT}\text{, and}$$

$$(16) \qquad \Delta(UNEMP/POP)_{jT} = \alpha + \beta_3 \hat{\Delta}(EMP/POP)_{jT} + \Delta X_{jT}\beta_4 + \delta_T + \varepsilon_{jT}\text{,}$$

where $\hat{\Delta}(EMP/POP)_{jT}$ is the instrumented change in state level employment.

Before performing these estimates, we test the relationship between projected demand shocks and state level employment to population rates over 1978 – 1998. Estimates are found in Table 6. The first row indicates that a one percentage point projected employment shock raises the employment to population ratio of high school dropout males by almost exactly one percentage point. For females, the employment impact is highly significant and about 60 percent as large. Inclusion of state dummies in the models reinforces these patterns.

Panel B presents corresponding estimates for those with high school or greater education. Not

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awards are not actually decided until a subsequent year.

surprisingly, adverse demand shocks have a larger impact on the employment of the less skilled.[42] Given these findings and the estimates above demonstrating the differential impact of DI recipiency on labor force participation for low skilled workers (Tables 3 and 4), we focus our inquiry on high school dropouts. Two-stage least squares estimates of (15) and (16) are found in Table 7.[43]

Panel A presents results for male high school dropouts. During 1979 – 84, a one percentage point decline in the employment to population ratio of male high school dropouts induced a 0.46 percentage point increase in non-participation rate and 0.54 percentage point increase in unemployment.[44] During the 1984 – 98, however, plausibly exogenous employment losses for high school dropout males appear to have lead to essentially one-for-one increases in labor force non-participation. Subsequent columns pool data for 1979 – 98, adding an interaction between the employment loss measure and a post-1984 dummy variable. These estimates confirm a significant increase in the labor force exit propensity of displaced high school dropouts males. Adding state fixed effects to the model does not appreciably affect the results. Panel B tabulates analogous estimates for female high school dropouts. Although the standard errors are slightly larger, the qualitative pattern of results is identical. Hence, there appears to be have been a significant increase in the propensity of high school dropouts to exit the labor force after job loss.

We have performed a variety of checks on the robustness of these results including using observations spaced at different intervals and performed on different sub-periods of 1984 – 1998. These tests confirm the finding that labor demand shocks feed significantly more into labor force exit post-1984 than was the case prior to DI liberalization. It bears emphasis, however, that these estimates do not imply that beginning in 1984, *all* high school dropouts who lose jobs exit the labor force. It is almost certainly the case that some displaced high school dropouts find reemployment immediately and hence are not

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[42] It is likely that low-skilled workers are employed in industries and occupations more vulnerable to shocks. In addition, their skills may be less mobile across sectors.

[43] Corresponding first-stage estimates are given in Appendix Table 1.

[44] All variables are denominated by the state population of male high school dropouts. The growth in the unemployment rate in conventional usage (i.e., denominated by labor force) would be approximately 30 to 50 percent larger.

captured by our demand shock instrument. Additionally, it is likely that the shock measure does not operate exclusively through job displacement. Adverse shocks may encourage some workers to exit the labor force voluntarily and, perhaps more importantly, slow the prevailing rate of reentry of non-participants into the labor force. In a similar vein, some share of those exiting may be previously unemployed workers whose re-employment prospects were harmed by the entry of new unemployed. Each of these forces will reduce the employment to population ratio and yield corresponding increases in non-participation. What appears unambiguous, however, is that state unemployment rates have become significantly less sensitive to adverse demand shocks while labor force non-participation and disability application rates have become significantly more so.

A few simple calculations demonstrate that the increased generosity of the DI program could plausibly account for much of this behavioral change. The coefficient estimates presented in the first three columns of Table 5 suggest that a one unit shock induced approximately a 0.12 percentage point greater increase in DI application rates in the post-1984 time period than prior to the 1984 reforms (e.g. 12 additional applications per 10,000 adults ages 25-64). Given that between 1984 and 1998, an average of 16 percent of the adult population aged 25-64 lacked a high school degree, we initially calculate that if all individuals induced to apply for DI were high school dropouts, then the fraction applying for DI in response to a one unit shock increased by 0.75 percentage points (0.0012/0.16).

This number should be compared with the estimates from Table 7, which suggest that the fraction of high school dropouts leaving the labor force in response to a one unit shock increased by 0.5 percentage points. Thus, the increased sensitivity of DI applications to adverse labor demand shocks in the post-1984 period could explain 150 percent of the change (0.75 / 0.50). If one makes the more reasonable assumption that only half of the individuals induced to apply for DI are high school dropouts, then the increase in DI application propensity can account for 75% of the increase. While this calculation is admittedly quite rough, it suggests that the DI program may be largely responsible for the increased propensity of low-skilled individuals to leave the labor force in response to job loss or declining labor market opportunities.

5. Alternative explanations: falling wages, immigration and incarceration.

In this section, we briefly explore three plausible alternative interpretations of the cross-state patterns of labor force withdrawal that are the focus of our study. One is declining real wages. A second is immigration of low skilled workers. And a third is the growth in the U.S. prison population.

a. Declining real wages

As argued by Juhn (1992) and Juhn, Murphy and Topel (1991), a stable elasticity of labor force participation coupled with declining real wages could explain the substantial decline in labor force participation of low skilled males over this period. To be clear, we view falling wages as a complement rather than a substitute to the role for the disability program in reducing labor force participation over 1984 – 1998; indeed, a key reason why disability replacement rates have risen is that wages at the lower tail of the distribution have fallen. For falling real wages to be the *primary* explanation for our findings, however, it would have to be the case that – not implausibly – wages fell by substantially more in states that experienced the most adverse demand shocks.

To explore the relevance of this hypothesis, Table 8 presents levels and changes of male high school dropout labor force participation and log real wages at five-year intervals over 1979 – 98. To facilitate regional comparisons, we divide the 50 states into three groups according to the magnitude of the demand shocks they are projected to have experienced over this 19 year interval ($\hat{\eta}_{79-98}$): the 10 most negative, the 10 least negative, and the remaining 30.[45] Consistent with well-known patterns, the data in Table 8

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[45] Details of our wage calculations are given in the Data Appendix. We focus on trends in male labor force participation since they are less likely to be driven by shifting gender norms. Because our demand shock measure, $\hat{\eta}_{79-98}$, is defined in *relative* terms, we convert it to an absolute measure to facilitate interpretation of the table. Specifically, we add an estimate of the overall demand shift experienced by high school dropouts in each five year interval equal to:

$$\overline{\eta}_{HSD,T} = \sum_k \phi_{kt}\gamma_{kt} \cdot \eta_{kT} ,$$

where ϕ_{kt} is the share of employees in industry (k) in year (t) who are high school dropouts, γ_{kt} is the industry (k) share of national employment in (t), and η_{kT} is the log change in the share of national employment between years (t) and (T). Unlike the *relative* shock measure in (12), this index need *not* be approximately mean zero in each period. Absolute demand shocks against high school dropouts are estimated at –0.048, -0.031, -0.034, and –0.019 log points in 79-84, 84-89, 89-94, and 94-98 respectively. Note that this transformation simply adds a constant to indicate that real wages and labor force participation of less-skilled males fell substantially during 1979 – 1998. We estimate that despite four years of strong growth during 1994 – 1998, male high school dropout wages declined by 22 log points and labor force participation by 5 percentage points over the 1979 – 1998 period.

Comparing wages and participation among the three groups of states arrayed in Table 8 reveals two key facts. First, labor force participation of high school dropout males fell by substantially more in the most adversely relative to the least adversely shocked states over 1979 – 1998: 11.0 percentage points in the former relative to 1.5 percentage points in the latter. Second, the differential patterns of labor force exit visible across these groups of states are *not* mirrored by differential wage declines. In fact, estimated log wages fell by somewhat less in the most adversely relative to the least adversely shocked states: 19 versus 25 log points. Hence, it is quite unlikely that differential regional wage declines alone can explain why labor force participation of male high school dropouts fell by substantially more in some regions than others.

In this vein, the growing availability and generosity of disability benefits appear a plausible alternative. Notably, the growth in DI recipiency per population was 1.1 percentage points in the most adversely shocked states over 1979 – 1998, relative to only 0.14 percentage points in the least adversely shocked states. Using the OLS estimates of the impact of DI recipiency on male high school dropout labor force participation found in Table 3, this difference can potentially explain a 5.0 to 7.8 percentage point relative decline in male high school dropout labor force participation between these two groups of states – 50 and 80 percent of the observed difference.[46] Hence, we tentatively conclude that while declining real wages played a crucial role in inducing labor force exit over 1984 – 1998, the disability system substantially exacerbated this trend by providing many of the least skilled with a viable alternative to

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each demand shock entry in Table 8 and does not impact the substantive results.

[46] The calculations are $(1.1 - 0.14)*(-5.18) = 5.0$, and $(1.1 - 0.14)*(-8.16) = 7.8$. Instrumental variables estimates of these same models that use levels and changes of state level replacement rates to identify exogenous variation in the change in DI recipiency in each state produce comparable point estimates. Note that the net change in disability per population over 1979 – 1998 understates the post-1984 growth since it also includes the pre-1984 contraction.

employment.

b. Cross-state patterns of immigration

The U.S. high school dropout population is increasingly composed of immigrants. Current Population survey data indicate that in 1998, 35 percent of non-elderly high school dropouts were foreign born, relative to only 12 percent of high school completers. Since legal immigrants are ineligible for disability benefits until five years after arrival, increased immigration flows would not appear to offer an alternative explanation for our primary findings. However, the flow of immigrants is quite geographically concentrated, with states such as New York and California receiving a disproportionate share of immigrants relative to population. If for any reason the high immigration states over 1984 – 1998 were also the 'low employment shock' states, this coincidence would induce a spurious positive cross-state correlation between state level adverse demand shocks and the labor force exit of high school dropouts.[47]

To explore the relevance of this alternative hypothesis, we reanalyzed our main results excluding the 11 states where foreign-born residents compose at least 40 percent of the high school dropout population.[48] Estimates that exclude these states continue to find that the share of male high school dropouts exiting the labor force in response to an adverse demand shock more than doubled during 1984 – 1998 relative to 1979 – 1984. Moreover, if we re-compute the Table 8 regional wage and employment comparison while dropping the 11 high immigration states, we again find that wages of high school dropouts fell roughly proportionately across all U.S. states while disability recipiency and labor force non-participation increased substantially more in the most adversely shocked states.[49] These results

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[47] For example, New York and California saw relatively little decline in high school dropout participation and relatively modest growth in disability recipiency over 1984 to 1999. Our analysis suggests that this is due to relatively favorable demand conditions prevailing in these states. In contrast, the immigration story implies that both facts are explained by the comparatively low level of DI eligibility among high school dropouts in these high immigration states.

[48] These states are AZ, CA, HI, FL, IL, MA, NJ, NV, NY, RI, and TX. In 1998, the foreign born share of high school dropouts in these states was 58 percent, relative to 11 percent in the remaining 39 states. Because the CPS does not provide information on country of birth prior to 1994, we are unable to exclude foreign-born high school dropouts from our sample.

[49] In fact, the regional wage results excluding the high immigrant states indicate a *more* uniform pattern of wage declines than is visible in Table 8. It is therefore likely that the relatively larger wage declines in 'low shock' states

suggest that differential cross-state patterns of immigration are not responsible for our main results.

c. Increased incarceration of high school dropouts

A third alternative we have considered for the declining participation of high school dropouts is their growing rates of incarceration (Freeman, 1991; Katz and Krueger, 1999; Western and Pettit, 2000). Because the incarcerated population is included in neither the numerator nor denominator of our labor force calculations, it is not intrinsically a source of bias for our participation estimates. It is likely the case, however, that potential criminals have below average rates of labor force participation relative to other high school dropouts. If it were also true that the most adversely shocked states experienced differentially high rates of incarceration growth over 1984 – 1998, this would mechanically raise measured labor force participation in these states.

To explore this possibility, we used data from the Bureau of Justice Statistics to measure incarceration rates of male high school dropouts across the three groups of states in Table 8. We find little difference in the share of non-elderly male high school dropouts in prison across these groups. For example, in our most adversely shocked states, the share of male high school dropouts incarcerated rose from 1.4 to 5.2 percent during 1984 – 98. In the least shocked states, it rose from 2.4 to 6.3 percent. The fact that we find comparable patterns of labor force exit for female high school dropouts – who have very low rates of incarceration – also suggests that the rising incarceration rate is not responsible for our results.

6. Conclusion: Implications for aggregate unemployment

Because high school dropouts were in 1984 differentially employed in industries that contracted during the next fourteen years, the reduced sensitivity of their unemployment to adverse labor demand shocks suggests that the share unemployed at present is substantially lower than it would have been in the absence of the DI program. Consider that the typical high school dropout in 1984 was employed in an

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over 1984 – 1998 are driven in part by a preponderance of immigrant high school dropouts in these states. Supplemental tables containing these results are available from the authors on request.

industry that over the next 14 years declined by 8.4 percentage points as a share of aggregate employment.[50] Combining this negative relative demand contraction with the estimated reduction in the unemployment responsiveness of high school dropouts from Table 7, we calculate that the share of high school dropouts who are presently unemployed would be 4.2 percentage points higher $(0.84 * 0.50 * 100 = 4.20)$ but for the liberalization of DI that occurred in 1984. Given that high school dropouts accounted for 12.2 percent of the non-elderly adult population in 1998, this suggests that the aggregate unemployment rate would have been approximately 0.64 percentage points higher in 1998 were it not for DI liberalization.[51]

To test the plausibility of this counterfactual, we take an entirely different approach to estimating it. At three-year intervals over 1984 – 1998, we calculate the sum of absolute employment declines in 2-digit industries nationally and weight these declines by the high school dropout share of employment in each. Specifically, we estimate:

$$(17) \qquad \hat{\delta}_{HSD,84-98} = \sum_t \sum_k \phi_{k,t} \min[Emp_{k,t} - Emp_{k,t+3}, 0],$$

where $\min[Emp_{k,t} - Emp_{k,t+3}, 0]$ is the absolute employment *loss* in industry k over 3 years and $\phi_{k,t}$ is the high school dropout employment share in industry k at the start of the interval. Using (17), we estimate cumulative job losses to high school dropouts totaling 791,000 between 1984 and 1998. Applying the Table 7 estimate of a 0.50 percentage point reduction over 1984 – 98 in the unemployment responsiveness of high school dropouts to unit demand shocks, we calculate that aggregate unemployment would have been approximately 0.30 percentage points higher in 1998.

This calculation is conservative in two ways. First, since most employment losses do not involve sustained national industry employment contractions over three years (consider that total U.S. employment grew by more than one third between 1984 and 1998), this estimate it is likely to detect only

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[50] To make this calculation, we estimate $\overline{\eta}_{HSD,T}$ for 1984 – 1998. See footnote 45.

[51] Because the unemployment rate is denominated by labor force and not by population, the estimated impact using a

a minority of all job losses. The estimate would be far higher, for example, if instead we counted all instances where an industry lost absolute employment at the *state* level over three years. Second, our calculation assumes that high school dropouts are no more likely than better-educated workers to lose work in the face of an employment contraction. In reality, job loss is much more commonplace among the less educated (Farber, 1997). If we amend (17) to assume that high school dropouts are two to three times as likely as better educated workers to lose work in a downturn, we estimate high school dropout job losses of 1.3 to 1.6 million workers, yielding a counterfactual 1998 aggregate unemployment rate that is 0.48 to 0.61 percentage points higher.

Reinforcing this calculation is the remarkable recent growth in the high school dropout DI recipient population. In 1999, almost 60 percent of all non-elderly male adult high school dropouts who were labor force non-participants were receiving disability benefits.[52] Given that the population of non-elderly high school dropouts declined by 30 percent between 1984 and 1998, we would have predicted a *contraction* of 550,000 in the high school dropout recipient population in these 14 years (making no allowance for the simultaneous decline in age and plausibly improved health of high school dropouts overall). Instead, the number of high school dropout DI recipients rose from 1.93 to 2.70 million. If even half of this unanticipated growth of 1.3 million high school dropout recipients is accounted for by the interaction of adverse shocks and increased program generosity, the unemployment rate would presently be one-half of a percentage point higher in the absence of programmatic changes.

Of course, some part of the change in the sensitivity of unemployment to adverse labor demand shocks may be driven by other factors – and thus our estimates could potentially be biased upwards. However, because the DI program has likely impacted labor force participation for at least some subset of the 88 percent of the labor force that is *not* composed of high school dropouts, it seems plausible that our estimate of the overall impact on the U.S. unemployment rate may understate the total effect.

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participation rate of 0.80 is 100*(0.042 * 0.122)/0.80 = 0.64.

[52] For female high school dropouts, the corresponding figure is 29 percent.

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Data Appendix

A.1 Employment, Unemployment, and Labor Force Non-Participation

Annual, state-level data on employment, unemployment and labor force non-participation by gender, age, and education category for individuals between the ages of 25 and 64 were calculated using the complete Current Population Survey monthly files for years 1978 – 1999. The number of observations ranges from 1.1 to 1.3 million annually. All calculations use CPS sampling weights. To attain comparable educational categories (high school dropout, high school graduate, some college, college-plus graduate) across the redefinition of Census's Bureau's education variable introduced in the 1992 CPS, we use the method proposed by Jaeger (1997). In particular, prior to 1992, we define high school dropouts as those with fewer than 12 years of completed schooling. In 1993 forward, we define high school dropouts as those without a high school diploma or GED certificate.

A.2. Wage Data from the CPS Merged Outgoing Rotation Group files

To estimate market wages for high school dropout males by state requires data on potential earnings for both workers and non-workers. Since our CPS samples do not provide earnings for non-participants, we impute them as follows using the CPS Merged Outgoing Rotation Group files for the years 1979 through 1999. For workers employed in the survey reference week, wages were calculated by dividing usual weekly earnings by usual weekly hours, multiplying top-coded earnings observations by 1.5, and inflating them to 1999 dollars using the chain weighted PCE deflator. We then estimated a regression in each year of the real log weekly wages of employed high school dropout males on a quartic in age, a set of state indicator variables, a dummy equal to one for black race, and a complete set of interactions between the black race dummy, the state indicators and the age quartic. Employing these regression coefficients, we assigned each unemployed or non-participant sample member the mean predicted wage of workers with identical characteristics, adding to their wage the 10th percentile residual from the relevant annual wage regression. Self-employed workers, who do not report earnings in most CPS samples, were assigned wages at the 50th percentile of the residual distribution. Averaging over observed and imputed wages, we calculate the mean earnings potential of high school dropouts in each state and year. In addition, we experimented with assigning non-participants the 25th and 50th percentile of the residual distribution. While these choices did affect the *level* of wages imputed in each year, the cross-state time pattern of results over 1979 – 1998 found in Table 8 is quite robust to the imputation method.

A.3. Simulated earnings histories from the CPS March annual demographic files

To estimate the deciles of the age-specific earnings profiles used in our replacement rate calculations, we utilize the Annual Demographic Files of the March Current Population Survey for the years 1964 – 1998 (n~50,000 per year). We include in our sample all males ages 25 – 64 who had positive earnings in the previous year and had positive hours in wage/salary employment in the survey reference week. Further details on these samples are provided in Katz and Autor (1999). We multiply weekly earnings deciles by (52/12) to estimate monthly earnings deciles for the PIA formula. To simulate earnings profiles, we calculate mean earnings at each percentile of the wage distribution within each state, year, age cell. The simulated earnings profiles assign to workers in a given state, age, percentile cell the earnings history of a worker who had mean earnings at the same percentile of the relevant state-age-year earnings distribution in each of the previous 15 years. Since our data do not allow us to calculate earnings histories beyond 16 years for workers observed in 1979 (the first year of the imputation), we assume that at years beyond this range, workers experienced wage growth equal to the mean wage of the economy of a whole. This assumption guarantees that imputed replacement rates neither rise nor fall in these years. We experimented with imputing longer earnings histories for workers observed later in our sample (e.g, up to 34 years of history for a worker of wage 59 observed in 1998). The level of benefits implied by these longer earnings histories was typically within 1 or 2 percent of the benefit level implied by our truncated earnings history simulation method.

Replacement rates based upon current earnings are calculated using the PIA formula in equation (2), setting the AIME equal to current monthly earnings at age specific percentiles of the monthly earnings distribution. Replacement rates based upon simulated earnings histories use equations (1) and (2). Data on the 'bend points' in the PIA formula were obtained from the Social Security Administration's *Annual Statistical Supplement.*

To test the plausibility of the increases in simulated replacement rates given in Table 2, we directly compared DI income and wage data for males by education group from Survey of Income and Program Participation data for 1984 and 1996. These comparisons, found in Appendix Table 2, indicate that the ratio of DI income to earnings for male high school dropouts rose from 42 to 54 percent between 1984 and 1996. Since this calculation compares DI income for the disabled to wage income for the currently employed, it is likely to substantially understate the true replacement rate. As has been documented elsewhere, disability recipients generally command below average earnings given education and experience prior to obtaining disability benefits (Bound, 1989).

A.4. DI and SSI Recipient and Benefits Data

Annual, state-level data on DI and SSI recipients, benefit levels, demographics and qualifying impairments were obtained from various years (1978 – 2000) of the Social Security Administration's *Annual Statistical Supplement.* DI data includes only disabled workers receiving benefits whereas SSI data includes only disabled adult beneficiaries (thus excluding child and aged beneficiaries).

A.5. DI and SSI Application and Award Data

Administrative data on disability applications and awards by state for years 1979 – 1998 were generously provided to us by Kalman Rupp and David Stapleton (DI), Charles Scott (SSI), and Alan Shafer of the Social Security Administration. In our data set, disability awards our dated according to the year of application rather than the year of decision. Because many applications filed after 1997 are still pending or under appeal, our final years of data underestimate the ultimate award rate.

A.6. DI and SSI Data from the Survey of Income and Program Participation

Disability recipiency rates by education category, age, and gender were estimated using data from the 1984 and 1996 waves of the Survey of Income and Program Participation. A survey respondent was coded as a DI recipient if he/she "did receive income from Social Security for himself/herself in this month" and whose reason for receipt of Social Security was disability. An individual was classified as an SSI recipient if he/she "did receive any income from Supplemental Security Income for him/her self during the reference period."

A.7: Calculating simulated changes in state level DI replacement rates

As an alternative to using the *level* of the replacement rate in 1984 as an instrument for the subsequent change in disability recipiency, we also exploit cross-state variation in the *change* in the replacement rate. An ideal test of the impact of replacement rates on labor force participation would exploit changes in the potential replacement earnings for each potential worker as a function of his or her entire earnings history. This approach requires a measure of the potential earnings of all individuals, including those who are not working, and is thus infeasible. We take the following alternative approach. We calculate the projected replacement rate for each employed worker based on his or her age and percentile in the wage distribution in 1984. Then, we simulate how mean replacement rates in each state are likely to have evolved over 1984 – 1998 as a function both of the increasing effective generosity of DI benefits and the changing age distribution in each state. This approach assumes that, conditional on age, there is a stable relationship between a worker's state of residence and his or her position – or, more precisely, the position of her potential earnings – in the national earnings distribution. Although this is clearly an approximation, it is likely to be reasonable assuming that average state earnings do not dramatically diverge over this time.

Define λ^{84}_{jad} as the probability that an individual residing in state (j) of age (a) in the year 1984 is in decile (d) of the national earnings distribution. Note that since λ^{84}_{jad} is a discrete density function, if follows that $\sum_{d=1}^{10}\lambda^{84}_{jad}=1$. Define $R^{t}_{ad}\in[0,1]$ as the DI mean replacement rate of a person of given age and earnings decile in year (t). Define γ^{t}_{ja} as the population shares of individuals of given age and state in year (t). Finally, define ϕ^{84}_{ja} as the labor force participation rate of workers in state (j) of age (a) in 1984. Hence, the mean potential replacement rate of those *currently employed* in state (j) in 1984 is:

$$(20) \qquad \overline{R}^{84}_{j}=\sum_{a=1}^{A}\sum_{d=1}^{10}\gamma^{84}_{ja}\cdot\phi^{84}_{ja}\cdot\lambda^{84}_{jad}\cdot R^{84}_{ad}\,.$$

To simulate the expected state level replacement rate in subsequent years, we allow both the age-decile replacement rate and the age distribution of individuals in each state to vary over time while holding constant the state level correspondence between age and earnings decile and the baseline participation rates of each stage-age group. The age categories used for this calculation are 25– 29, 30 – 34, 35 – 39, 40 – 44, 45 – 50, 50 – 54, and 55 – 61. Hence, the simulated change in the mean state replacement rate for state (j) from 1984 to a later period (T) is equal to:

$$(21) \qquad \hat{\Delta}\overline{R}^{84-T}_{j}=\sum_{a=1}^{A}\sum_{d=1}^{10}\hat{\phi}^{84}_{ja}(\hat{\gamma}^{T}_{ja}-\hat{\gamma}^{84}_{ja})\cdot\hat{\lambda}^{84}_{jad}(\hat{R}^{T}_{ad}-\hat{R}^{84}_{ad})\,.$$

Note that because labor force participation rates of low wage workers dramatically declined over 1984 – 1998, it would be potentially quite misleading to perform this calculation using the wages of *employed* workers in later years. Our approach abstracts from changes in labor force participation by simulating state replacement rates only as a function of the changing effective generosity of DI benefits and changes in the age distribution of state populations.

Figure 1: DI/SSI Recipients and DI Mortality Rate: 1978 - 1999

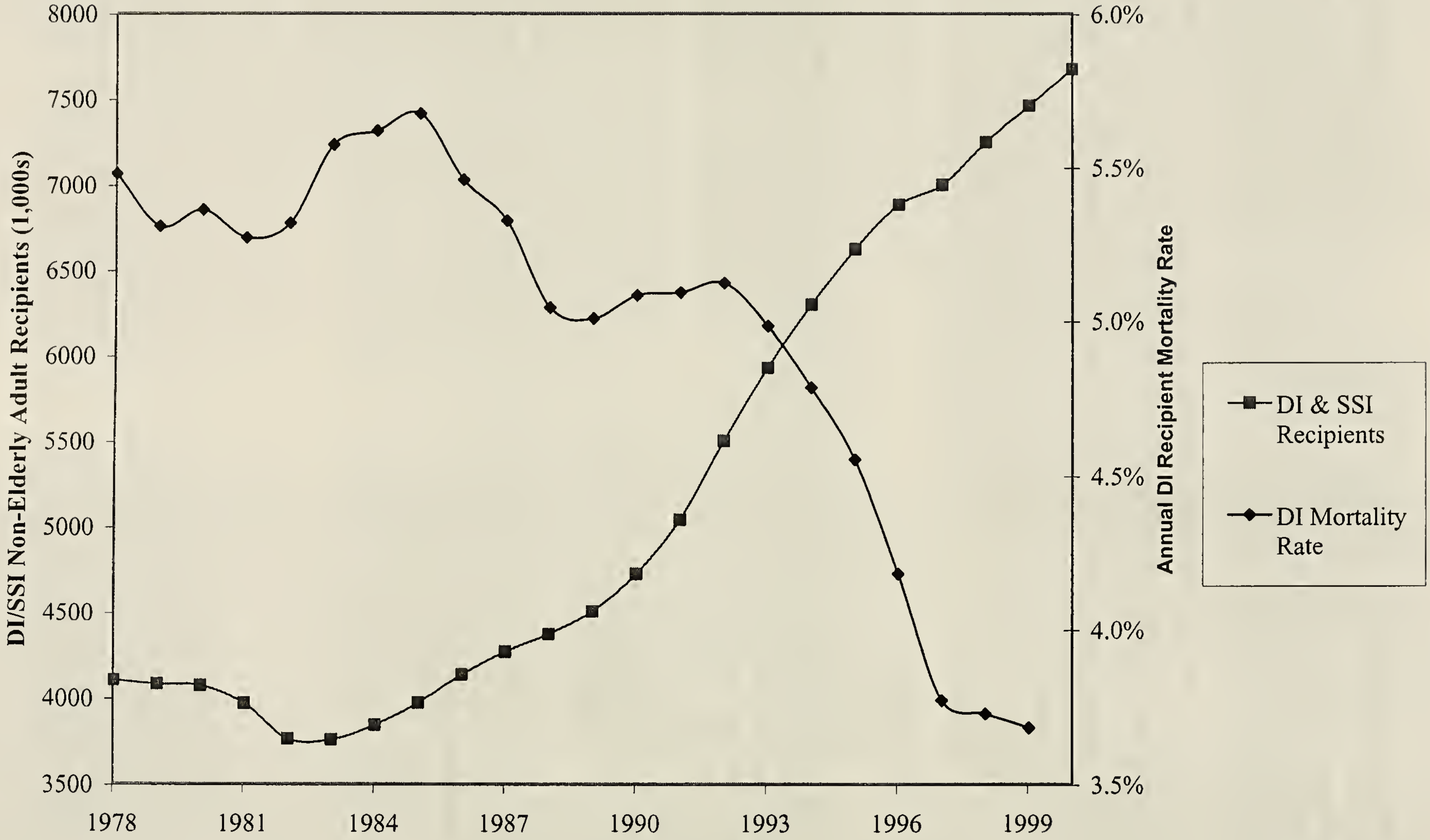


Figure 2: Disability Award Rate Per Population: 1979 - 1999 (1979 = 100)

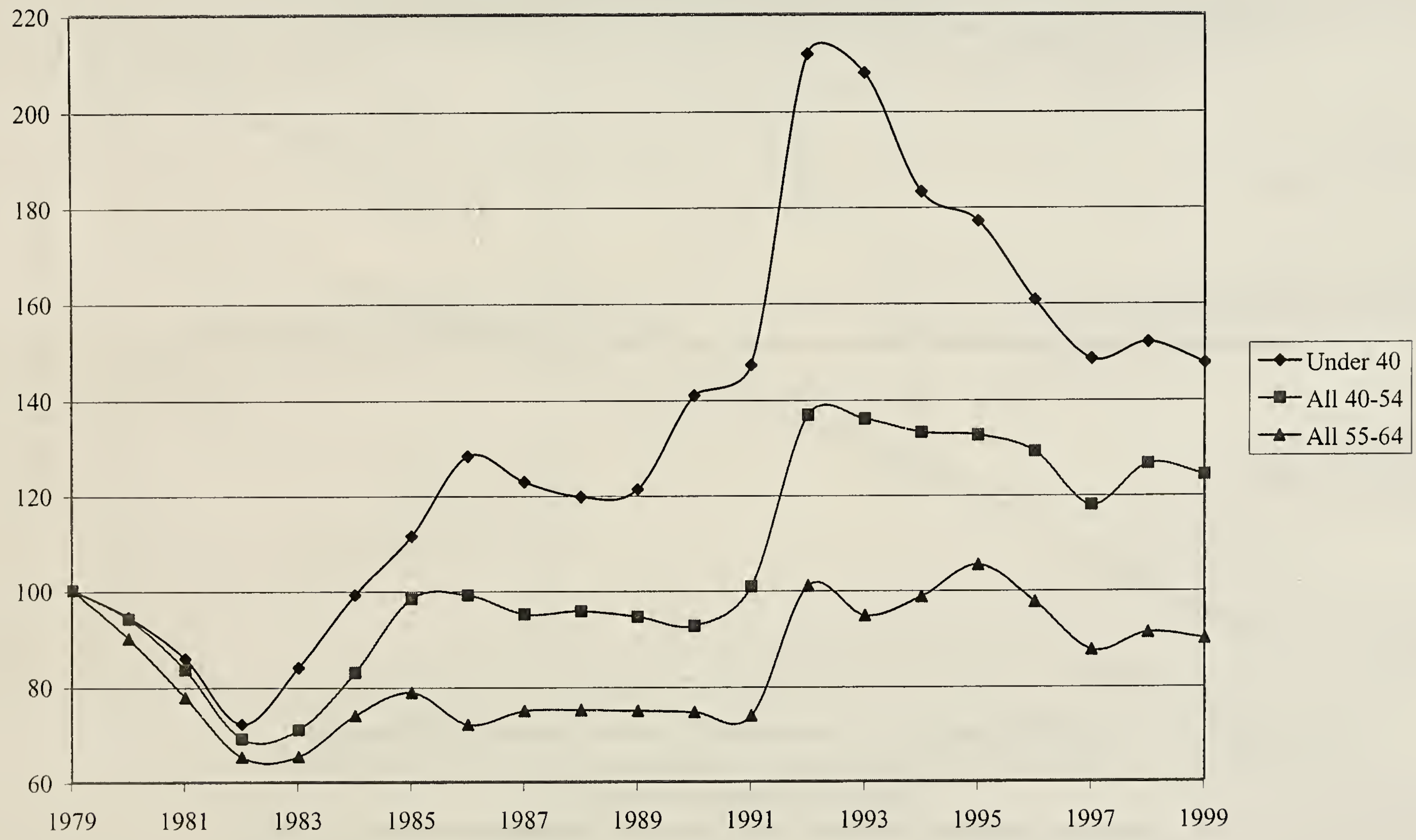
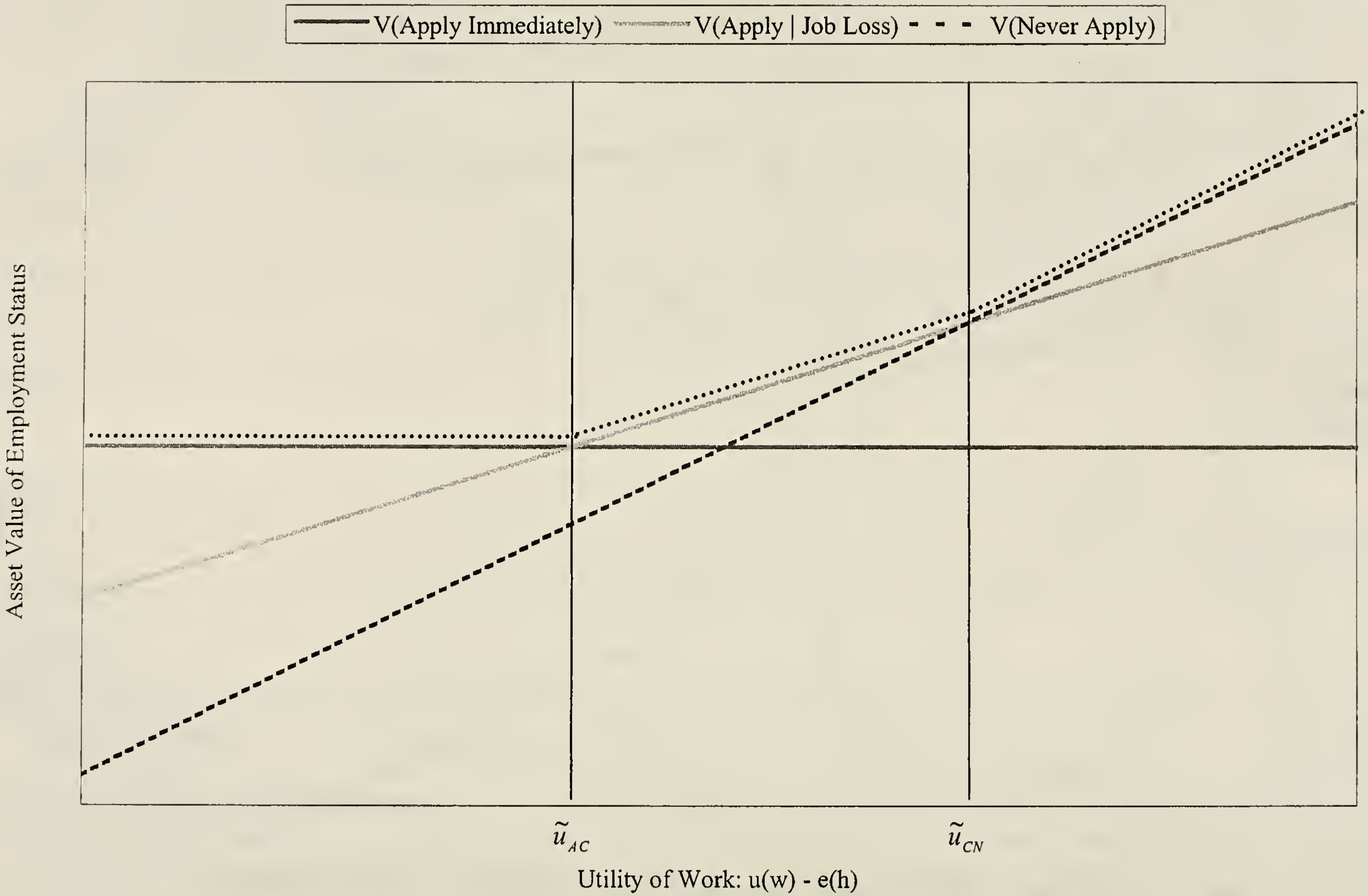


Figure 3: The Choice of Disability Applicant Status - Always, Conditional, and Never - as a Function of Earnings and Health



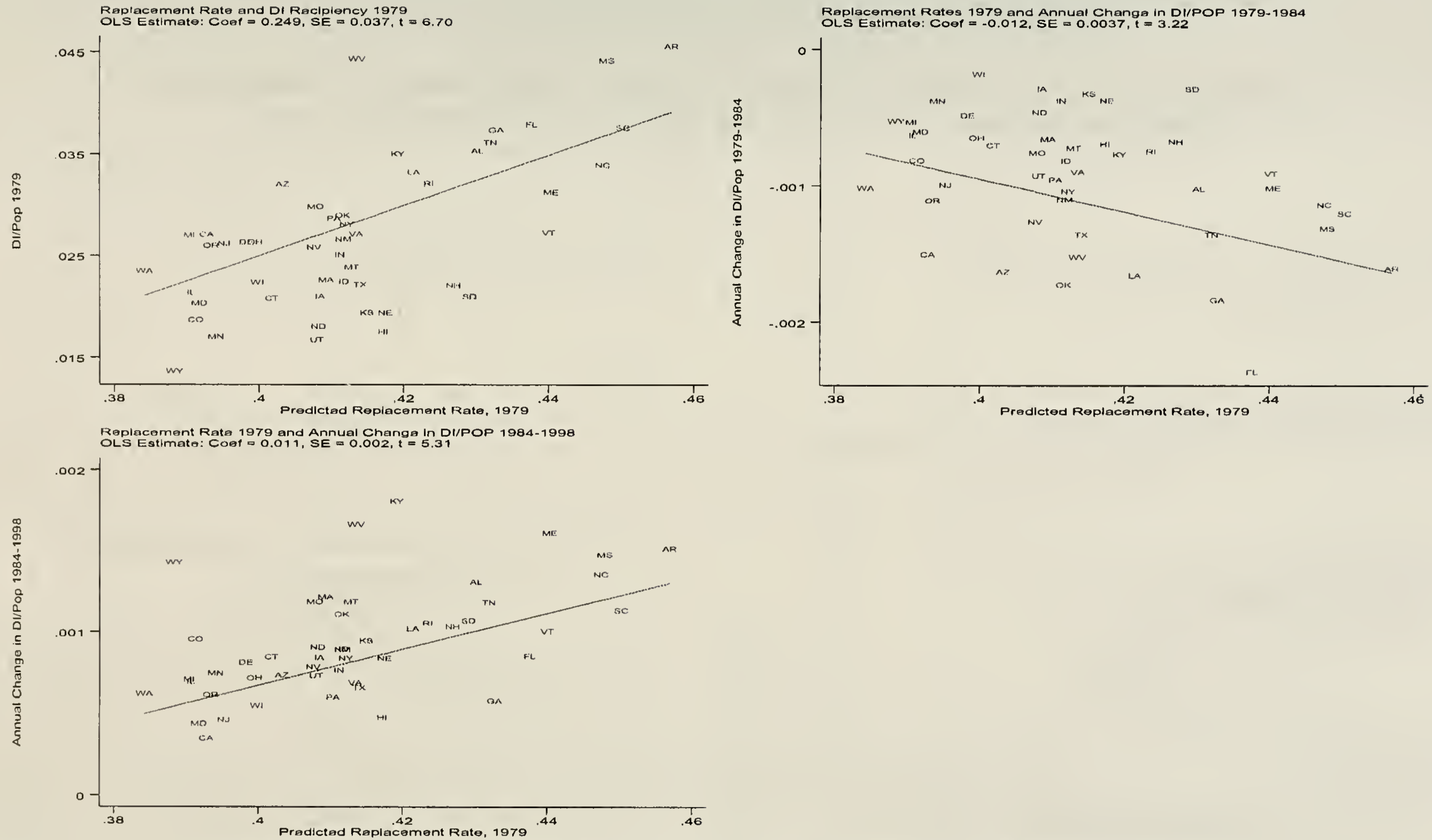


Figure 4: Simulated State DI Replacement Rates 1979 and Levels and Changes of DI Recipiency Per Adult Population, 1979-98

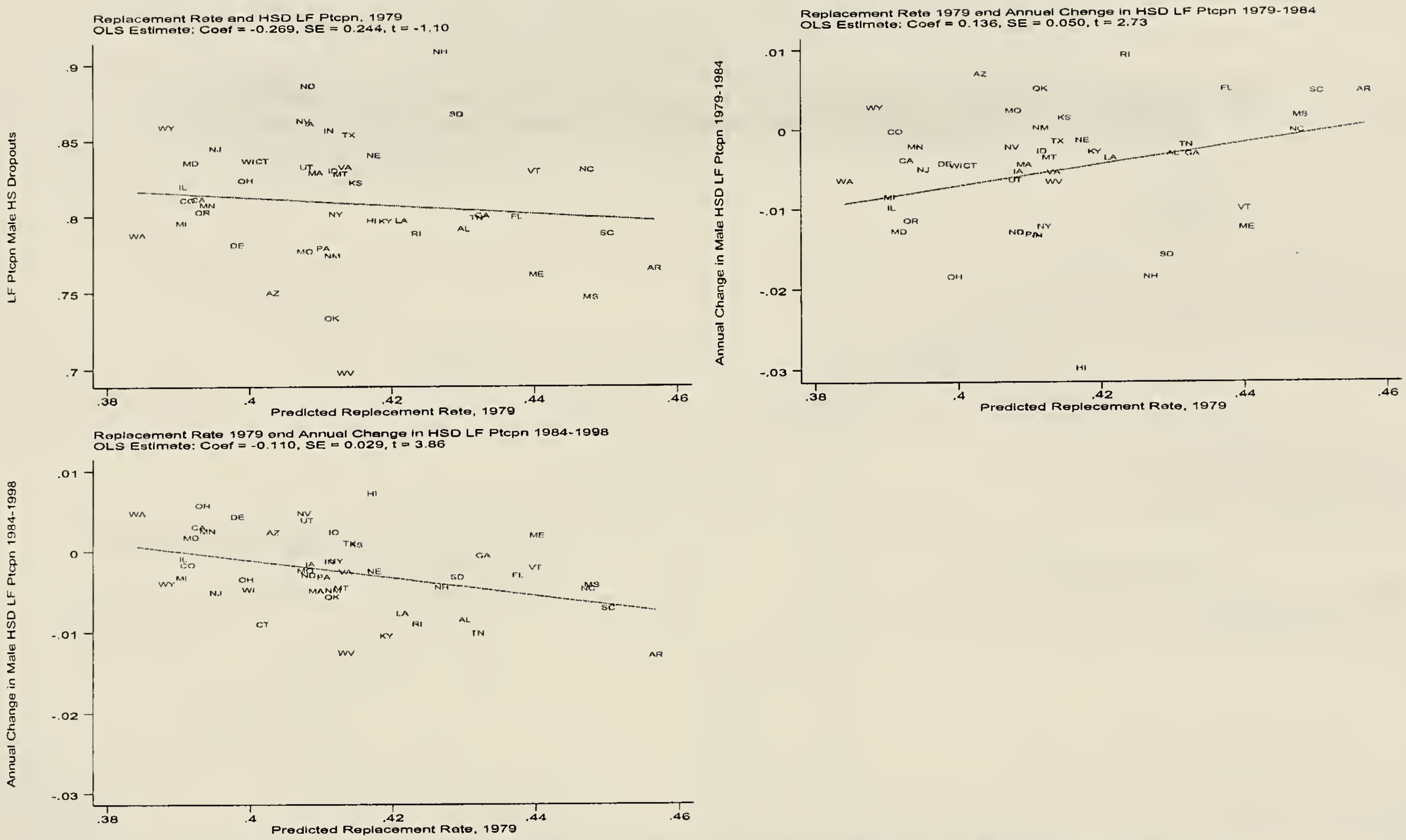


Figure 5: Simulated State DI Replacement Rate 1979 and Levels and
Changes of Male HSD Labor Force Participation, 1979-98

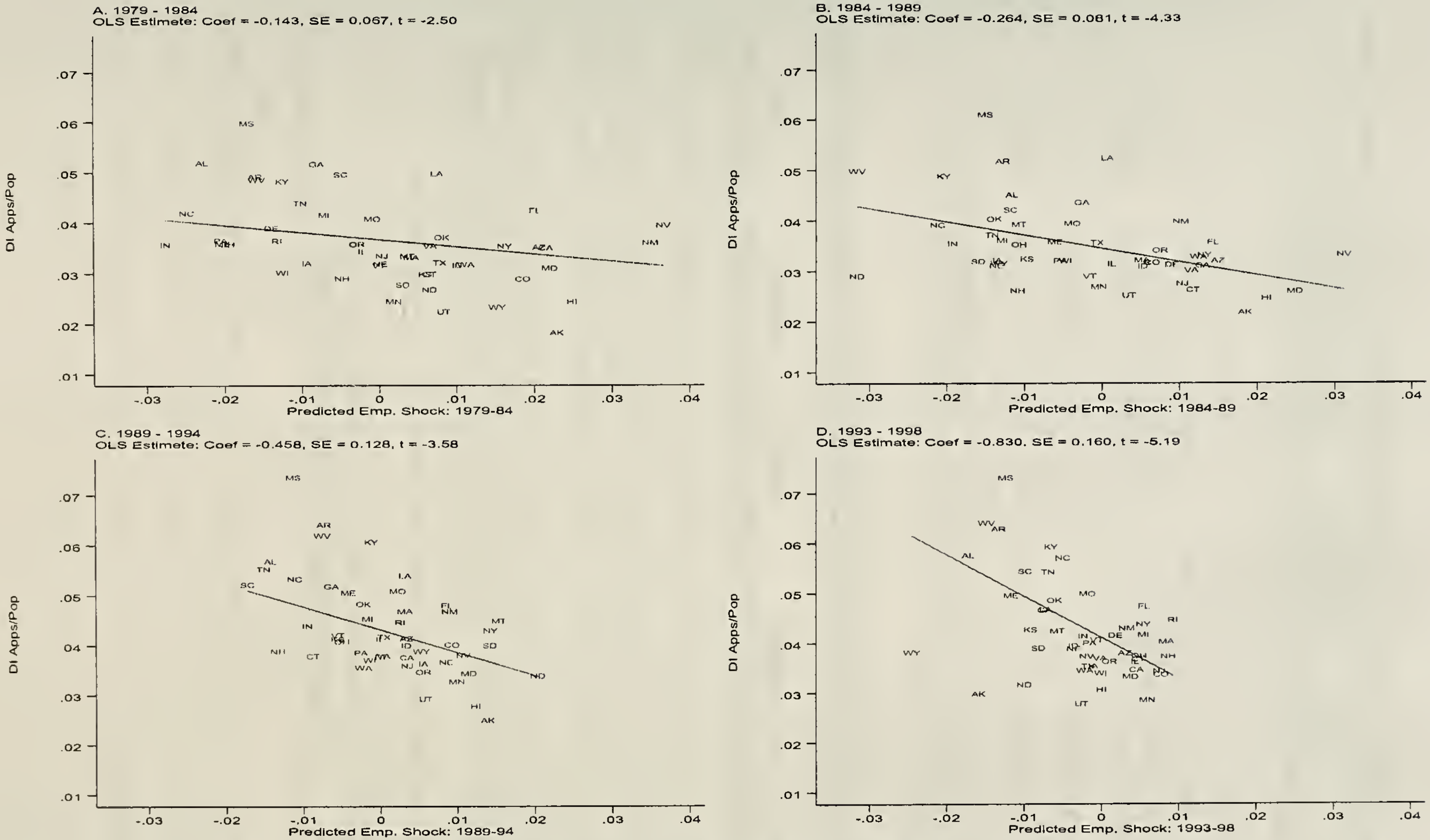


Figure 6: Predicted State Employment Shocks and

Disability Applications per Population, 1979 - 1998

Table 1A. Changes in the Health of DI Awardees, 1983 - 1998

Diagnosis Category	Four Year Mortality Rate	Percentage of DI Awards 1983	1988	1993	1998
Neoplasms	81%	17%	13%	13%	11%
Circulatory	20%	22%	18%	14%	13%
All others	16%	32%	32%	33%	32%
Mental disorders	5%	16%	21%	26%	22%
Musculo-skeletal	5%	13%	17%	15%	23%

Source: Social Security Administration, Annual Statistical Supplement, 1984, 1989, 1994 and 1999. Four-year mortality rate is from administrative follow-up of those awarded benefits in 1985.

Table 1B: Percent of Adults Ages 25 - 64 Receiving Either DI or SSI in 1984 and 1999

	25-39		*40-54*		*55-64*		*25-64*	
Education	**1984**	**1999**	**1984**	**1999**	**1984**	**1999**	**1984**	**1999**
HS Dropouts	3.5%	8.6%	7.4%	15.5%	13.9%	28.9%	8.0%	15.9%
HS Graduates	1.4%	4.1%	2.4%	5.5%	5.9%	10.4%	2.6%	5.7%
Some College	0.7%	1.4%	1.7%	3.6%	4.6%	7.3%	1.5%	3.1%
College Graduates	0.2%	0.4%	0.7%	1.3%	1.5%	3.3%	0.5%	1.2%
High School Plus	0.9%	2.1%	1.8%	3.5%	4.3%	7.5%	1.7%	3.4%

Source. Authors' tabulations, Survey of Income and Program Participation, 1984 and 1999.

Table 1C: Percent of Labor Force Non-Participants Ages 25 - 64 Receiving Either DI or SSI in 1984 and 1999

	25-39		*40-54*		*55-64*		*25-64*	
Education	**1984**	**1999**	**1984**	**1999**	**1984**	**1999**	**1984**	**1999**
HS Dropout Males	30%	47%	49%	61%	47%	60%	45%	57%
HS Dropout Females	6%	16%	16%	29%	18%	44%	14%	29%
HS Plus Males	19%	27%	40%	40%	21%	25%	25%	31%
HS Plus Females	3%	7%	5%	14%	7%	17%	4%	12%

Source. Authors' tabulations, Survey of Income and Program Participation, 1984 and 1999.

Table 2: Earnings Replacement Rates for Male Disability Recipients at Various Percentiles of the Wage Distribution: 1979 and 1998

A. Males Ages 50 - 54

Earnings Percentile	Current Earnings Only 1979	1989	1998	Simulated Earnings History 1979	1989	1998
10%	48.6%	50.6%	52.7%	53.3%	58.9%	63.3%
25%	43.3%	44.1%	45.0%	47.7%	50.7%	52.7%
50%	36.8%	37.7%	38.4%	38.7%	40.9%	42.6%
75%	31.3%	31.1%	31.9%	32.5%	32.9%	35.2%
90%	27.4%	26.7%	26.6%	28.4%	27.4%	28.0%

B. Replacement Rate at 10th Percentile

Ages	Current Earnings Only 1979	1989	1998	Simulated Earnings History 1979	1989	1998
55-61	49.9%	52.8%	56.3%	56.3%	65.2%	74.1%
50-54	48.6%	50.6%	52.7%	53.3%	58.9%	63.3%
40-49	48.4%	50.3%	52.9%	52.4%	56.7%	57.4%
30-39	48.5%	53.0%	56.9%	50.1%	55.4%	59.8%

Replacement rates are calculated using Social Security Administration Disability Insurance benefit formula for 1979 and 1998 in conjunction with weekly earnings data from March CPS files for 1964 - 1998. See text for details.

Table 3: State DI and SSI Recipiency and Labor Force Participation of Males and Females Ages 25 - 64, 1979 - 1984 and 1984 - 1998

A. Δ Male Labor Force Participation

	High School Dropouts				High School Grad Plus			
	1979-1984		*1984-1998*		*1979-1984*		*1984-1998*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ SSI & DI Recipiency	-3.23		-3.12		-0.33		-0.55	
	(0.99)		(0.71)		(0.36)		(0.21)	
Δ DI Recipiency		-8.16		-5.18		-0.67		-0.14
		(2.01)		(1.55)		(0.76)		(0.50)
Δ SSI Recipiency		4.26		-1.51		0.21		-0.89
		(2.87)		(1.29)		(1.12)		(.42)
R^2	0.37	0.46	0.58	0.60	0.18	0.19	0.19	0.21

B. Δ Female Labor Force Participation

	High School Dropouts				High School Grad Plus			
	1979-1984		*1984-1998*		*1979-1984*		*1984-1998*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ SSI & DI Recipiency	-1.00		-1.17		1.08		0.20	
	(1.20)		(0.80)		(0.60)		(0.44)	
Δ DI Recipiency		-4.07		-5.68		0.73		-0.48
		(2.66)		(1.98)		(1.37)		(1.01)
Δ SSI Recipiency		3.98		2.47		1.59		0.74
		(4.04)		(1.66)		(1.88)		(0.85)
R^2	0.07	0.10	0.21	0.31	0.18	0.18	0.02	0.03

n =50 U.S. states. Standard errors in parentheses. Estimates also control for change in the age distribution in the state population (ages 25 - 39 and 40 - 54 with 55 - 64 omitted) for each relevant education group. Estimates are weighted by mean state share of U.S. population in the two years to form the dependent variable. Weighted mean of Δ SSI & DI recipiency/population is -0.00597 from 1979-1984 and 0.02247 from 1984-1998.

Table 4: OLS and Instrumental Variables Estimates of the Relationship Between DI Applications and Labor Force Participation of Males and Females Ages 25 - 64: 1979 - 1984 and 1984 - 1998

A. Δ Male Labor Force Participation

	Male High School Dropouts					Male High School Grad Plus				
	1979-84	*1979-84*	*1984-98*	*1984-98*	*1984-98*	*1979-84*	*1979-84*	*1984-98*	*1984-98*	*1984-98*
	OLS	**IV**	**OLS**	**IV**	**IV**	**OLS**	**IV**	**OLS**	**IV**	**IV**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10**
Δ DI Apps/ Pop	-1.66 (0.83)	-5.23 (2.26)	-1.61 (0.47)	-1.69 (0.99)	-3.25 (1.16)	-0.25 (0.28)	-0.70 (0.57)	-0.01 (0.14)	0.26 (0.31)	0.25 (0.32)
R^2	0.28	.	0.53	0.53	0.40	0.18	0.14	0.07	.	0.01
Instrument		*Rep rate 1979*		*Rep rate 1984*	*ΔRep 1984-98*		*Rep rate 1979*		*Rep rate 1984*	*ΔRep 1984-98*
1st Stage Coef.		-0.025 (0.007)	. .	0.025 (0.007)	2.005 (0.578)		-0.028 (0.007)	. .	0.023 (0.006)	1.986 (0.561)

B. Δ Female Labor Force Participation

	Female High School Dropouts					Female High School Grad Plus				
	1979-84	*1979-84*	*1984-98*	*1984-98*	*1984-98*	*1979-84*	*1979-84*	*1984-98*	*1984-98*	*1984-98*
	OLS	**IV**	**OLS**	**IV**	**IV**	**OLS**	**IV**	**OLS**	**IV**	**IV**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10**
Δ DI Apps/ Pop	0.54 (0.95)	-3.89 (2.47)	-0.95 (0.53)	-1.36 (1.13)	-0.38 (1.21)	0.87 (0.50)	1.77 (1.20)	1.16 (0.67)	-0.13 (0.60)	0.64 (0.63)
R^2	0.06	.	0.23	0.22	0.21	0.18	.	0.18	0.00	.
Instrument		*Rep rate 1979*		*Rep rate 1984*	*ΔRep 1984-98*		*Rep rate 1979*		*Rep rate 1984*	*ΔRep 1984-98*
1st Stage Coef.		-0.027 (0.007)	. .	0.022 (0.006)	1.862 (0.556)		-0.025 (0.008)	. .	0.021 (0.006)	1.897 (0.536)

n =50 U.S. states. Standard errors in parentheses. Δ disability applications per population is constructed as the change in DI apps/pop between the start and end years of the period multiplied by the number of elapsed years. For 1979 - 1984, we add 1/2 to this number the estimated number of disability beneficiaries terminated per population for not meeting medical standards in each state over 1979 - 1984. Estimates also control for change in the age distribution in the state population (ages 25 - 39 and 40 - 54 with 55 - 64 omitted) for each relevant education group. Estimates are weighted by mean state share of U.S. population in the two years to form the dependent variable. Instruments are simulated state mean DI replacement rate 1979 (columns 2 & 7), simulated state mean DI replacement rate 1984 (columns 4 & 9), and change in simulated state mean replacement rate 1984 - 98 (columns 5 & 10). See text for details.

Table 5: Impacts of Projected Log Employment Shocks on DI Applications per Adult Population Ages 25 - 64, 1978 - 1998 at 3 Year Intervals

	A. DI Apps/Pop				B. DI Awards/Pop			
	1978-84	*1984-98*	*1978-98*	*1978-98*	*1978-84*	*1984-98*	*1978-98*	*1978-98*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
E[Δln(State Emp)]	-0.086	-0.212	-0.057	-0.002	-0.079	-0.123	-0.054	-0.003
	(0.039)	(0.040)	(0.036)	(0.018)	(0.020)	(0.024)	(0.021)	(0.013)
E[Δln(State Emp)]*			-0.173	-0.049			-0.083	-0.018
(1985 - 98 dummy)			(0.052)	(0.024)			(0.030)	(0.017)
State Dummies	No	No	No	Yes	No	No	No	Yes
R^2	0.36	0.44	0.40	0.90	0.43	0.44	0.46	0.86
n	100	250	350	350	100	250	350	350

Standard errors are in parentheses. Dependent variable in Panels A is cumulative (unique) DI applications per state population over three year intervals. Dependent in Panel B is DI awards per state population, calculated similarly. All models include year dummies and controls for the education, gender, and age composition of state populations in each year. Control variables measure changes in share of state population by gender, three education categories (high school dropout, high school grad, and some college, with college omitted), and two age categories (25-39 and 40-54, with 55-64 omitted) separately by gender.

Table 6: Estimated Impact of Projected Log Employment Shocks on Employment to Population Ratio of Males and Females Ages 25 - 64, 1978 - 1998 at Three Year Intervals

	A. Δ Emp/Pop High School Dropouts			
	Males		Females	
	(1)	(2)	(3)	(4)
E[Δln(State Emp)]	1.026	1.294	0.580	1.074
	(0.258)	(0.372)	(0.246)	(0.350)
State Dummies	No	Yes	No	Yes
R^2	0.45	0.48	0.31	0.37
	B. Δ Emp/Pop High School Plus			
	Males		**Females**	
	(1)	(2)	(3)	(4)
E[Δln(State Emp)]	0.250	0.363	0.085	0.263
	(0.091)	(0.133)	(0.109)	(0.153)
State Dummies	No	Yes	No	Yes
R^2	0.47	0.50	0.41	0.47

n=350 in each column. Standard errors in parentheses. All models include year dummies and control for state population shares age 25-39 and 55 -64 (40-54 omitted) for relevant education group. Estimates are weighted by state share of U.S. population ages 25 - 64 in each year.

Table 7: Instrumental Variables Estimates of the Impact of Employment Shocks on Labor Force Exit of High School Dropous Ages 25 - 64: 1979 - 1984 and 1984 - 1998

A. High School Dropout Males

	Δ NILF/Pop				Δ Unemp/Pop			
	1979-84	*1984-98*	*1979-98*	*1979-98*	*1979-84*	*1984-98*	*1979-98*	*1979-98*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ Emp/Pop	-0.462	-1.021	-0.558	-0.614	-0.538	0.021	-0.442	-0.386
	(0.170)	(0.187)	(0.193)	(0.220)	(0.170)	(0.187)	(0.193)	(0.220)
Δ Emp/Pop*			-0.435	-0.502			0.435	0.502
(85 - 98 dummy)			(0.232)	(0.289)			(0.232)	(0.289)
State Dummies	No	No	No	Yes	No	No	No	Yes
R^2	0.71	0.68	0.71	0.66	0.46	0.32	0.51	0.43
n	100	250	350	350	100	250	350	350

B. High School Dropout Females

	Δ NILF/Pop				Δ Unemp/Pop			
	1979-84	*1984-98*	*1979-98*	*1979-98*	*1979-84*	*1984-98*	*1979-98*	*1979-98*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ Emp/Pop	-0.489	-1.273	-0.641	-0.523	-0.511	0.273	-0.359	-0.477
	(0.273)	(0.378)	(0.215)	(0.207)	(0.273)	(0.378)	(0.215)	(0.207)
Δ Emp/Pop*			-0.616	-0.598			0.616	0.598
(85 - 98 dummy)			(0.363)	(0.375)			(0.363)	(0.375)
State Dummies	No	No	No	Yes	No	No	No	Yes
R^2	0.70	0.81	0.81	0.84	.	.	.	.
n	100	250	350	350	100	250	350	350

Standard errors in parentheses. All estimates include year dummies and control for changes in the age distribution in the state population (ages 25 - 39 and 40 - 54 with 55 - 64 omitted). Estimates are weighted by state share of U.S. population in each year. Instrument for change in employment to population ratio of high school dropouts is projected state level log employment shocks and (in columns 3, 4, 7 and 8) its interaction with a 1985 - 98 dummy. See text for details of construction of instruments.

Table 8: High School Dropout Labor Force Participation and Mean Log Wages of Male Ages 25 - 64 in High, Medium, and Low Projected Demand Shock States, 1979 - 1998.

	A. Most Negative Shocks		*B. Intermediate Shocks*		*C. Least Negative Shocks*	
	LF/ Pop	**Log Wage**	**LF/ Pop**	**Log Wage**	**LF/ Pop**	**Log Wage**
1979	0.798	2.352	0.817	2.418	0.802	2.394
	(0.044)	(0.144)	(0.028)	(0.126)	(0.018)	(0.096)
1984	0.764	2.244	0.789	2.316	0.775	2.289
	(0.053)	(0.124)	(0.035)	(0.117)	(0.035)	(0.088)
1989	0.738	2.172	0.766	2.230	0.792	2.246
	(0.055)	(0.116)	(0.052)	(0.135)	(0.030)	(0.079)
1994	0.683	2.124	0.725	2.132	0.753	2.081
	(0.068)	(0.093)	(0.060)	(0.117)	(0.061)	(0.087)
1998	0.688	2.164	0.751	2.198	0.788	2.144
	(0.078)	(0.112)	(0.064)	(0.096)	(0.050)	(0.076)

D. Projected Demand Shocks and Changes in Participation and Log Wages over 5 year intervals, 1979 - 1998

	LF/Pop	**Wage**	**η**	**LF/Pop**	**Wage**	**η**	**LF/Pop**	**Wage**	**η**
Δ 79 - 84	**-0.034**	**-0.108**	**-0.068**	**-0.028**	**-0.102**	**-0.049**	**-0.028**	**-0.105**	**-0.028**
Δ 84 - 89	**-0.026**	**-0.072**	**-0.045**	**-0.023**	**-0.086**	**-0.033**	**0.017**	**-0.043**	**-0.017**
Δ 89 - 94	**-0.055**	**-0.048**	**-0.040**	**-0.041**	**-0.098**	**-0.035**	**-0.039**	**-0.165**	**-0.027**
Δ 94 - 98	**0.006**	**0.041**	**-0.026**	**0.027**	**0.066**	**-0.019**	**0.034**	**0.063**	**-0.015**
Δ 79 - 98	**-0.110**	**-0.187**	**-0.179**	**-0.066**	**-0.220**	**-0.136**	**-0.015**	**-0.250**	**-0.086**

Cross-state standard deviations in parentheses. Most negatively shocked states are AL, AR, IA, IN, KY, MS, NC, PA, WV, WY. Least negatively shocked states are AK, AZ, CA, FL, HI, MD, NM, NV, NY, WA. Middle group includes all other U.S. states except DC. Wages for high school dropout males labor force non-participants are imputed at the 10th percentile of the residual earnings distribution. Demand shock measures (η) are calculated excluding own-state employment. See text for details.

Appendix Table 1: First Stage Estimates of the Impact of Projected State Employment Shocks on Employment to Population Ratio of High School Dropous Ages 25 - 64: 1979 - 1984 and 1984 - 1998

	Models Estimated by Period				Pooled Models with Period Interactions			
	Males		*Females*		*Males*		*Females*	
	Δ Emp/Pop		Δ Emp/Pop		Δ Emp/Pop		Δ Emp/Pop	
	1979-84	1984-98	1979-84	1984-98	*1979-84	*1985-98	*1979-84	*1985-98
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
E[Δln(State Emp)]	0.799	1.175	0.585	0.489	1.014	-0.057	1.006	0.167
	(0.297)	(0.377)	(0.267)	(0.356)	(0.386)	(0.338)	(0.362)	(0.326)
E[Δln(State Emp)]*					0.608	1.353	-0.006	0.642
(1985-1998 dummy)					(0.532)	(0.465)	(0.501)	(0.452)
Δ Pop Share 25-39	0.002	0.006	-0.012	-0.210	0.000	-0.013	-0.184	-0.190
	(0.159)	(0.079)	(0.131)	(0.070)	(0.077)	(0.067)	(0.065)	(0.059)
Δ Pop Share 55-64	-0.200	-0.174	-0.279	-0.414	-0.168	-0.125	-0.394	-0.327
	(0.132)	(0.088)	(0.119)	(0.071)	(0.080)	(0.070)	(0.065)	(0.058)
State Dummies	No	No	No	No	Yes	Yes	Yes	Yes
R^2	0.21	0.40	0.16	0.36	0.49	0.44	0.39	0.40
n	100	250	100	250	350	350	350	350

Standard errors in parentheses. All estimates include year dummies and are weighted by state share of U.S. population in each year. Age shares refer to state high school dropout population in the relevant gender. Omitted age group is 40 - 54. See text for details of construction of instruments.

Appendix Table 2: Lower Bound Estimates of DI Earnings Replacement Rates for Males by Education Group from Survey of Income and Program Participation: 1984 and 1996

	Mean Monthly Earnings	DI Income	After Tax Replacement Rate
		1984	
High School Dropouts	$ 1,001	$ 390	41.9%
High School Grads	$ 1,209	$ 436	38.8%
Some College	$ 1,392	$ 468	36.2%
College Plus	$ 1,835	$ 534	31.3%
		1996	
High School Dropouts	$ 1,216	$ 609	54.2%
High School Grads	$ 1,704	$ 668	42.4%
Some College	$ 2,047	$ 682	36.1%
College Plus	$ 3,240	$ 810	27.1%

Eamings data are for all employed males from SIPP 1984 wave 1 month 1 and SIPP 1996 wave 1 month 1. DI income data is for all males receiving Disability from the same samples. Replacement rates account for 7 percent payroll tax in 1984 and 7.65 percent payroll tax in 1996 (not paid on DI income).