THE RISE IN THE DISABILITY ROLLS AND THE DECLINE IN UNEMPLOYMENT*

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Between 1984 and 2001, the share of nonelderly adults receiving Social Security Disability Insurance income (DI) rose by 60 percent to 5.3 million beneficiaries. Rapid program growth despite improving aggregate health appears to be explained by reduced screening stringency, declining demand for less skilled workers, and an unforeseen increase in the earnings replacement rate. We estimate that the sum of these forces doubled the labor force exit propensity of displaced high school dropouts after 1984, lowering measured U. S. unemployment by one-half a percentage point. Steady state calculations augur a further 40 percent increase in the rate of DI receipt.

The federal Disability Insurance (DI) program is the largest income replacement program in the United States directed toward nonelderly adults, with annual cash transfers exceeding $54 billion in 2001. Once benefits are awarded, recipients receive income replacement and health insurance through Medicare until return to work, medical recovery, death, or retirement at age 65, at which point they obtain equivalent benefits from the Social Security Administration's Old Age and Survivors Insurance program. Over the past two decades, two important changes have impacted eligibility and generosity for this program. The first, occurring in 1984, liberalized the disability determination process, reversing a dramatic reduction in disability rolls underway. The second occurred more gradually. Because the DI benefits formula is progressive and indexed to the mean wage in the economy, the widening dispersion of earnings during the 1980s and 1990s substantially raised the ratio of disability in-

come to prior earnings (the "replacement rate") for low-skilled workers, a trend augmented by the rising real value of medical benefits provided. Contemporaneously, the share of nonelderly adults receiving DI income rose by more than 60 percent from its 1984 trough to 3.7 percent of the adult population ages 25 to 64 (5.3 million beneficiaries). Relative to prior DI cohorts, new entrants were younger and substantially more likely to suffer from low mortality impairments, particularly pain and mental disorders.

The objective of this paper is to assess how these programmatic changes—reduced screening stringency and a rising replacement rate—interacted with declining demand for less-skilled workers to impact labor force participation of the low skilled during the period of 1978 to 1998. The question is intrinsically difficult. DI is for the most part a uniformly administered federal program, making for limited treatment-control comparisons. Additionally, while aggregate health improved during the past two decades, reduced DI screening stringency coincided with a period of declining labor market opportunities for less-skilled workers, making it likely that *demand* for benefits (as a form of income replacement) rose while benefits *supply* was shifting outward.[1] Finally, the administrative definition of disability—the inability to "engage in a substantial gainful activity"—ensures that holding health constant, disability benefits are easier to obtain when job opportunities are scarcer. Given these confounds, distinguishing the impacts of increasing supply from increasing demand for benefits is a major theme of our paper.

To (attempt to) surmount these issues, we develop two sets of instrumental variables to proxy demand and supply conditions. To identify exogenous variation in the supply of disability benefits, we exploit the progressivity of the DI benefits formula. This formula does not account for variation in regional wage levels, and hence potential replacement rates are substantially higher in low wage states. This variation yields differential cross-state benefit supply shifts during program contractions and expansions (both of which occurred in our sample frame) in high versus low replacement rate states. To identify exogenous variation in the demand for benefits, we follow Bartik [1991] in constructing a measure of state level labor demand shifts at low and medium frequency. This measure, a weighted sum of national industry

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1. See Juhn, Murphy, and Topel [1991], Juhn [1992], and Katz and Autor [1999] on adverse demand shifts faced by low-skilled workers in recent decades.

employment changes (excluding own state employment) projected onto state industry composition, performs well in predicting the flow of DI applications by state, indicating that adverse demand shocks raise demand for disability benefits. Exploiting this combination of instruments, we study the impact of the supply and demand for DI benefits on the labor force behavior of low-skilled workers during the period of 1978–1998.

Our contribution builds on an influential literature exploring the impacts of disability benefits on the level of labor force participation in the United States and abroad.[2] Distinct from this literature, our paper analyzes the impact of the supply of disability benefits on the responses of low-skilled workers to adverse labor market shocks. The rationale for this focus is that nonemployment is a de facto precondition for disability application, making the opportunity cost of seeking DI benefits higher for employed workers than for job losers and nonparticipants. Many potential beneficiaries will therefore seek benefits in the event of job loss but not otherwise. We call these potential beneficiaries "conditional applicants." As our conceptual model demonstrates, the size of this conditional group is likely to be relatively elastic to program and labor market conditions. By contrast, direct employment to disability transitions will be relatively inelastic. Our framework therefore implies that reduced DI screening stringency, rising replacement rates, and slackening demand for low-skilled workers should have raised the propensity of workers facing an adverse demand shock to exit the labor force to seek disability benefits.

This prediction structures our empirical analysis. After establishing that cross-state shifts in the supply of benefits raised the labor force participation of less-skilled workers during the 1979–1984 disability retrenchment and lowered them thereafter,

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2. Key papers in this literature include Parsons [1980], Bound [1989], Haveman, De Jong, and Wolfe [1991], Bound and Waidman [1992], Gruber and Kubik [1997], 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. Lewin-VHI, Inc. [1995], Rupp and Stapleton [1995], and Stapleton et al. [1998] analyze the importance of the economic climate to disability application and receipt. Bound and Burkhauser [1999] provide a comprehensive overview of the labor market impacts of disability programs. The work closest in spirit to the approach taken here is the excellent study by Black, Daniel, and Sanders [2002], who measure the impact of shocks to coal prices on DI and SSI income in mining intensive counties. The Black, Daniel, and Sanders study does not explore labor market implications.

we turn to an examination of the interaction of the disability program with adverse demand shocks. We find that DI application rates for given demand shocks rose secularly after the disability reforms of 1984, reaching two to three times their prereform levels by the late 1990s. Paralleling this change, we estimate a 60 percent increase in the propensity of displaced high school dropouts to exit the labor force. Simple calculations suggest that increased disability application propensity can account for this behavioral change. As alternative explanations for the observed labor force exit of less-skilled workers, we explore the importance of changing mortality rates, rising immigration and incarceration rates, fluctuations in UI benefits, and, logically, falling wages. Although many of these factors appear relevant to cross-state patterns of low-skilled labor force participation, none substantially alters the finding that the increasing supply of DI benefits induced substantial labor force exit of low-skilled workers during 1984–1998.

To gauge the importance of changes in the DI program for the aggregate labor market, we calculate the demand contraction experienced by high school dropouts over the post-disability-reform years to form a counterfactual labor force participation figure net of disability. A limitation of our approach is that it does not allow us to disaggregate the separate labor market impacts of reduced screening and rising replacement rates—the latter a partial function of declining relative wages for low-skilled workers. Summing over these structural components, our "reduced-form" counterfactual suggests that the changes in labor force behavior of less-skilled workers induced by the disability system lowered the measured U. S. unemployment rate of nonelderly adults by half a percentage point since 1984. Notably, the impact of disability on the aggregate labor market is yet to be fully felt. Steady state calculations augur a further 40 percent increase in DI recipiency rates over the next decade, which is likely to be concentrated among less-skilled workers.

I. POST-1984 CHANGES TO THE FEDERAL DISABILITY SYSTEM: SCREENING, BENEFITS, AND BENEFICIARIES

The federal government provides cash and medical benefits to the disabled through the Social Security Disability Insurance

(DI) and Supplemental Security Income (SSI) programs.[3] The health eligibility criteria for the two programs are identical, requiring a medically determinable impairment that prevents the applicant from engaging in any "substantial gainful activity." SSI benefits are means-tested and do not require prior work history. DI benefits, which are not means tested, are set according to a recipient's prior earnings history. To obtain benefits, applicants provide detailed medical, income, and asset information to a federal Social Security Administration (SSA) office, which makes the disability determination. Individuals currently in the labor force are not normally eligible for disability benefits.[4]

I. A. Clampdown and Liberalization

During the late 1970s, concern over swelling disability rolls spurred the Social Security Administration to tighten medical eligibility criteria and exercise greater control over the state boards that interpret SSA's eligibility standards. The fraction of applicants awarded benefits (the "award rate") fell from 45 percent in 1976 to 32 percent in 1980. Augmenting this administrative action, Congress passed legislation in 1980 mandating that SSA conduct more frequent beneficiary health reassessments known as Continuing Disability Reviews. In the subsequent three years, SSA determined that more than 380,000 DI beneficiaries—40 percent of those whose cases were reviewed—no longer met medical standards and terminated their benefits [Rupp and Scott 1998]. Congress also required the Social Security Administration to further tighten medical criteria, accelerating the decline in award rates. This large-scale curtailment of benefits, occurring during the deepest postwar U. S. recession, was met with intense public criticism. Citing violations of due process, seventeen states refused to comply with the DI review effort during 1983 and 1984.

Responding to the backlash, Congress passed legislation in 1984 that profoundly altered the disability determination system,

3. Approximately one-fourth of DI recipients also receive funds from Supplemental Security Insurance (SSI), which is an entitlement program rather than a labor income insurance system. Though the two programs have many overlaps, we focus on DI because by design, it has far more interaction with labor force participants. Bound, Burkhauser, and Nichols [2001] report that approximately 85 percent of DI applicants were employed 36 months prior to application while the comparable figure for SSI applicants was below 30 percent.

4. For example, earnings exceeding $500 per month in 1999 would have automatically disqualified a DI applicant.

yielding a broader definition of disability and providing appli-
cants and medical providers with greater opportunity to influence
the decision process.[5] Despite improving economic conditions, the
number of DI awards increased by one-third from its 1982 trough
to a 1986 peak (the highest level reached during the 1980s).
Contemporaneously, Continuing Disability Reviews came to a
near halt. In the five years from 1985 through 1989, SSA termi-
nated fewer individuals for failing to meet medical eligibility
standards than it had terminated in the first five months of 1982.

I. B. Rising Replacement Rates

While we have stressed the role of the liberalization of
screening in expanding the supply of DI benefits, potentially as
significant was an unforeseen rise in the earnings replacement
rate. This rise was caused by the interaction between the DI
benefits indexation schedule and the growth of earnings inequal-
ity in the United States economy.[6]

Determination of an individual's DI benefit proceeds in two
steps. First, the beneficiary's Average Indexed Monthly Earnings
(AIME) is computed as

$$(1) \qquad AIME_i = \frac{1}{T} \sum_{t=1}^{T} Y_{it} \cdot \max\left[\frac{\bar{Y}_{T-2}}{\bar{Y}_t}, 1\right],$$

where Y_{it} is equal to an individual's average monthly earnings
(conditional on employment) in each year t, inflated to current

5. 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, such as ability to function in a work or worklike setting; 2)
consider source evidence provided by the applicant's own health care provider
prior to the results of SSA consultative examination; 3) give additional weight to
pain and related factors; 4) consider multiple nonsevere impairments as consti-
tuting a disability during the initial determination (whereas prior to 1984, appli-
cants were automatically denied awards during the initial determination if all
impairments were judged to be nonsevere); 5) desist from terminating benefits for
any individual for whom SSA could not demonstrate substantial evidence of
medical improvement; 6) provide benefits for those former recipients whose ter-
minations were under appeal; and 7) suspend Continuing Disability Reviews
(CDRs) for mental impairments and pain until appropriate guidelines could be
developed. In the post-1984 period, two additional administrative factors affected
applications and terminations. In 1991, due to successful court challenges to
SSA's treatment of source evidence, regulations were adopted placing further
weight on the information provided by an SSI or DI applicant's own medical
provider. Finally, agency downsizing during the 1980s and increased claims
workload in the 1990s resulted in a substantial decrease in the frequency of CDRs
during 1989–1993. See Stapleton et al. [1998, pp. 66–69] for a detailed discussion.
6. To our knowledge, this increase has been overlooked by the economics and
DI policy literatures.

dollars by the ratio of the average wage in the United States economy two year's prior ($\bar{Y}_{T-2}$) to the average wage in the year of earnings.[7]

Second, the benefit awarded, the "Primary Insurance Amount" (PIA), is computed from the AIME using the piecewise linear formula,

(2)

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

where the "bend points" ($b1$, $b2$) are also rescaled each year by average wage growth in the economy. As visible in (2), the DI benefits formula is progressive (i.e., concave), ensuring that low-wage workers replace a greater share of income.[8] Indexation of the benefit formula to the mean wage in the economy further ensures that benefit levels keep pace with aggregate earnings growth.

In an era of stable earnings inequality, this formula has little impact on the evolution of replacement rates. Between 1979 and 1995, however, real weekly earnings of full-time, full-year workers with less than a high school degree fell by 19.1 percentage points, while the Social Security Administration's mean wage series increased by 21.6 percentage points in real wage terms.[9] This increase in the proportional difference between mean wages and below-mean wages caused the DI replacement rate for low-wage workers to rise substantially.

To illustrate the mechanism underlying this rise, it is useful to consider a worker whose earnings growth over her career lagged contemporaneous average nominal wage growth (as would be true for many low-skilled workers in this era). Because the

7. In addition, approximately the lowest five years of earnings are discarded and quarters with earnings in excess of that period's taxable Social Security maximum are truncated at the cap. Once the DI benefit is awarded, annual cost-of-living increases are tied to the Consumer Price Index. The two-year lag for the earnings indexation factor in equation (1) reflects the historic time lag required for calculating the numerator of this series.

8. For example, a worker with an AIME that did not exceed $b1$ would receive 90 percent earnings replacement. Because a 7.65 percent Social Security payroll tax is assessed on wage but not on DI income, the effective replacement rate would be closer to 100 percent.

9. Real wages are calculated from CPS March Annual Demographic files as annual earnings of full-time male high school dropouts ages 25–64 for earnings years 1979 and 1998. Wages series are deflated by the chain-weighted PCE deflator.

wage index applied in (1) and (2) inflates prior earnings to current
dollars using the growth rate of average nominal wages, the
indexed monthly earnings (AIME) computed in step 1 of the
benefits calculation for this low earnings-growth individual would
exceed current earnings. Additionally, because the bend points in
(2) follow the same wage index, these would also have risen
relative to individual earnings. A greater share of the worker's
AIME would therefore be replaced on the steeper sections of the
curve (i.e., in the 90 and 32 percent ranges rather than the 15
percent range). These two impacts—a rising AIME and rising
bend points—are additive: indexation of earnings to the mean
wage raises the ratio of AIME to prior earnings while indexation
of the bend points raises the share of the AIME replaced on the
more generous sections of the benefits formula.

The distributional impacts of indexation are seen in Table I.
Currently employed male workers ages 55–61 at the tenth per-
centile of the (age-specific) earnings distribution who obtained DI
benefits in 1979 would have replaced 52 percent of their current
earnings with DI cash transfers. By 1999, this number had risen
to 74 percent. Accounting for the rising value of in-kind Medicare
benefits, the potential DI replacement rate for a tenth percentile
male age 55–61 rose still further from 67 to 104 percent. Nor was
this rise limited to the older workers. For males at the tenth
percentile of earnings, the rise in the replacement rate exceeded
18 percentage points in all age brackets. At higher positions in
the earnings distribution, however, the growth was far less pro-
nounced. For workers at the seventy-fifth and ninetieth percen-
tiles of earnings, potential replacement rates rose a compara-
tively modest 4 to 8 percentage points.[10]

I. C. Changing Characteristics of DI Beneficiaries

As documented in Tables II and III, several marked demo-
graphic shifts in the DI population accompanied these benefit

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10. Further information on all data sources and variable construction is
provided in the Data Appendix. Note that there was also a substantial—albeit far
less pronounced—rise in the DI replacement rate for relatively high wage workers
due to a large increase in earnings subject to the OASDI tax. Income subject to the
OASDI tax was capped at approximately 1.2 times average earnings during the
1950s, 1960s, and early 1970s and then rose to approximately 2.4 times average
by 1989, where it currently remains. Because benefits are only computed from
taxed rather than total earnings, the increase in the cap caused replacement rates
to rise for even relatively high wage workers.

TABLE I
POTENTIAL DI INCOME AS A PERCENTAGE OF CURRENT EARNINGS FOR NONELDERLY
MALES AT VARIOUS PERCENTILES OF THE WAGE DISTRIBUTION, 1979 AND 1999

Age	Earnings percentile	Cash income replacement rate		Adding in-kind Medicare benefit	
		1979	1999	1979	1999
55–61	10	52	74	67	104
	25	45	54	48	63
	50	37	47	36	47
	75	27	32	26	31
	90	20	24	19	23
50–54	10	47	57	61	81
	25	41	47	44	55
	50	34	41	33	42
	75	26	32	25	31
	90	19	23	18	22
40–49	10	48	53	61	80
	25	41	45	44	55
	50	34	41	33	42
	75	26	33	25	32
	90	20	26	19	25
30–39	10	46	54	59	84
	25	41	46	44	58
	50	36	41	35	44
	75	29	36	27	35
	90	23	28	21	27

Replacement rates are calculated using the Social Security Administration Disability Insurance benefit formula for 1979 and 1999 in conjunction with annual earnings data from the March CPS files for the years 1963–1998 and the 1978 CPS Social Security Earnings Records Exact Match file for the years 1951–1962. The first two columns represent the ratio of potential disability benefits to current earnings for males in the labor force and with nonzero earnings in 1978 and 1998. The latter two columns add average Medicare expenditures to DI benefits and average percentile-specific fringe benefits to earnings to estimate a total compensation measure of the replacement rate. To calculate Average Indexed Monthly Earnings (AIME) for an M year old Nth percentile worker in year T, we use equation (1) of the text and assume that this individual was an $M - t$ year old Nth percentile worker in year $T - t$ and that low earnings years (those excluded in the AIME calculation) occurred before the age of 25. Percentile ranks in this calculation are year and age specific. We use equation (2) of the text to estimate potential DI benefits as a function of the AIME. The final two columns add average Medicare expenditures for DI beneficiaries in the relevant year to the DI benefit amount and scale earnings to account for average fringe benefit rates for individuals at each of the five earnings percentiles using data from Pierce [2001]. See the Data Appendix for further details.

supply shifts. One was a rapid increase in the share of younger recipients. Among males ages 25–39 and 40–54, DI receipt rose by 50 percent between 1984 and 1999. The proportional increase for males above age 54 was only one-quarter as large. Among women, growth in DI receipt was even more rapid but also skewed toward younger recipients. On net, the share of DI re-

TABLE II

DI RECEIPT AND LABOR FORCE PARTICIPATION BY GENDER, EDUCATION, AND AGE
1979, 1984, AND 1999

	A. Males									B. Females								
	All			HS dropout			HS plus			All			HS dropout			HS plus		
Age	79	84	99	79	84	99	79	84	99	79	84	99	79	84	99	79	84	99
A. DI Recipients per 1000 nonelderly adults (SSA and Survey of Income and Program Participation data)																		
25–39	11	10	15		21	53		5	11	4	4	10		7	21		4	8
40–54	35	28	42		52	105		18	26	15	12	30		35	60		10	21
55–64	113	96	108		148	201		46	59	51	43	72		92	164		29	62
B. Percent of nonelderly, nonparticipants receiving DI benefits (Survey of Income and Program Participation data)																		
25–39		17.2	17.0		23.5	26.8		14.1	13.7		1.3	3.2		1.5	3.9		1.1	3.0
40–54		36.9	32.7		38.5	40.0		35.6	29.6		4.9	10.0		7.4	11.7		3.3	9.5
55–64		30.2	26.6		42.5	43.2		20.3	20.3		8.7	16.8		14.2	24.4		5.1	13.8
C. Percent of nonelderly adults participating in labor force (Current Population Survey data)																		
25–39	95.7	94.7	93.1	91.0	88.1	86.1	96.6	95.8	94.1	63.9	70.0	76.3	49.6	50.3	55.0	66.9	73.2	78.7
40–54	92.7	92.7	90.2	86.5	85.0	76.3	95.4	95.0	91.9	60.3	65.7	77.4	48.8	49.8	54.0	64.9	70.6	80.1
55–64	73.0	68.6	68.1	64.2	60.2	53.2	79.0	73.3	71.2	41.9	41.8	51.6	33.8	33.3	32.4	47.0	46.0	55.7
D. Percent of nonelderly adults unemployed (Current Population Survey data)																		
25–39	3.7	6.3	3.1	7.0	12.5	6.0	3.0	5.4	2.7	3.9	5.0	3.2	6.2	8.3	6.7	3.4	4.5	2.8
40–54	2.5	4.3	2.4	4.0	7.2	4.5	1.9	3.4	2.2	2.4	3.4	2.1	3.1	4.8	3.8	2.2	2.9	1.9
55–64	1.9	3.4	1.9	2.3	4.7	2.1	1.7	2.6	1.8	1.3	1.8	1.4	1.5	2.4	1.7	1.1	1.4	1.3

DI receipt rates by age and gender in Panel A are calculated using recipient counts from Social Security Administration, *Annual Statistical Supplement* [various years] denominated by population estimates from the Current Population Survey Merged Monthly files. Education-specific DI receipt in Panel A and the fraction of labor-force nonparticipants receiving DI in Panel B are calculated with the Survey of Income and Program Participation, 1984 wave 1 and 1996 wave 12. Panels C and D were calculated from Current Population Survey Merged Monthly files.

cipients between the ages of 40 and 54 rose by more than 50 percent following the 1984 liberalization while the share of male beneficiaries declined by 15 percent [U. S. Social Security Administration, *Social Security Bulletin: Annual Statistical Supplement,* various years].

The DI population has always been substantially less educated than average, and as the program grew post-1984, it encompassed a substantially larger share of this population. Using data from the Survey of Income and Program Participation (SIPP), we estimate that the share of high school dropouts receiv-

TABLE III
DISTRIBUTION OF QUALIFYING IMPAIRMENTS OF DI AWARDEES AT FIVE-YEAR
INTERVALS, 1983–1999

Qualifying impairment	4-Year mortality rate (%)	Percent of DI awards			
		1983	1988	1993	1999
Neoplasms	81.0	16.8	13.2	12.6	10.6
Circulatory disorders	19.8	21.9	17.6	14.0	12.1
Musculo-skeletal disorders	5.3	13.4	16.8	14.8	23.7
Mental disorders	5.4	16.3	20.9	26.1	22.5
All others	16.0	31.6	31.5	32.5	31.1

Source: Social Security Administration, *Annual Statistical Supplement,* 1984, 1989, 1994, and 2000. Four-year mortality rate is from administrative follow-up of those awarded benefits in 1985 [Hennessey and Dykacz 1993].

ing disability benefits rose by over 60 percent after 1984, and more than doubled for high school dropout males ages 40–54.[11] In 1999, a 40–54 year old high school dropout was four to five times as likely to receive DI benefits as a male in the same age range with at least a high school degree.
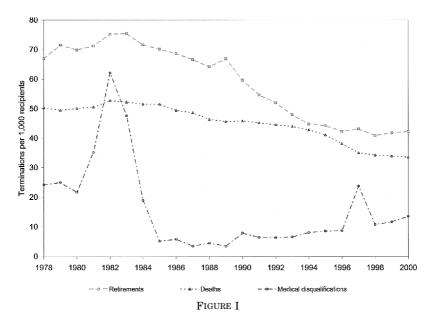
The rapid growth in DI receipt among male high school dropouts mirrored a substantial decline in their labor force participation (panel C). Between 1984 and 1999 the labor force participation rate of high school dropout males declined by 8.7 percentage points among those 40–54, and 7 percentage points among those 55–64. Simultaneously, the share of male high school dropouts in these age brackets receiving DI rose by 5.3 and 6.3 percentage points. Hence, despite the steep decline in male high school dropout participation, the share of male high school dropout *nonparticipants* receiving DI benefits in these age categories rose to slightly above 40 percent.[12]

Finally, as younger beneficiaries entered the DI rolls, the fraction suffering from comparatively low mortality impairments

11. Since SSA does not report the educational distribution of DI recipients, we use the Survey of Income and Program Participation (SIPP) to estimate these numbers for 1984 and 1999. SIPP data are unfortunately not available for earlier years. Relative to high school dropouts, proportionate growth in DI receipt for those with at least a high school degree was substantially smaller among males and slightly larger among females. In all cases, growth among better educated workers was from a low base.

12. Unemployment per population rates of high school dropouts also fell steeply during these years (panel D). We discuss these trends further in Section V.

FIGURE I

DI Termination Rates per 1000 Beneficiaries by Reason, 1978–2000

Source: Social Security Bulletin: Annual Statistical Supplement [various years]. Termination rates are equal to the fraction of DI beneficiaries terminated by cause annually.

grew (Table III). The share of DI awardees with a primary diagnosis of a mental disorder or a disease of the musculo-skeletal system (typically back pain)—the two disorders with the lowest mortality among SSA's fourteen major diagnostic categories—increased by 60 percent between 1983 and 1999. The corresponding award shares for neoplasms (cancers) and circulatory system diseases (primarily heart disease), both of which have mortality far in excess of average, declined by 40 percent.

Several consequences of these demographic shifts are seen in Figure I, which plots the annual rate of DI benefit termination by cause from 1978–2000: death, retirement, and medical disqualification. Following the 1984 liberalization, DI rolls increased at an average annual rate of 4.2 percent (*excluding* dependents of DI recipients). As younger cohorts with lower mortality impairments entered, the annual mortality rate of DI recipients fell by 35 percent, and the exit rate into the retirement system declined by 40 percent. Accordingly, the expected benefit duration of newer

cohorts substantially exceeds 1984 levels [Rupp and Scott 1998].[13]

Because most objective evidence suggests that the prevalence of disabling illness fell during this time [Cutler and Richardson 1997], it is likely that the growth in DI rolls is primarily explained by nonhealth factors that shifted the supply and demand for disability benefits. These include a more expansive definition of disability, changes in the DI award process, rising replacement rates and, closely related, declining labor market opportunities for less skilled workers.[14]

II. WHEN DO SHIFTS IN BENEFITS SUPPLY IMPACT LABOR SUPPLY: DEMAND SHOCKS OR DIRECT QUITS?

Before analyzing the impact of changing DI benefits supply and demand on the labor market, we offer a brief model to motivate our empirical approach. In a market setting where all job separations were involuntary, workers would apply for DI benefits only at the onset of illness or job loss. An analysis of the impact of shifts in the supply of benefits on labor force participation would therefore focus on the labor supply decisions of displaced workers, specifically whether they chose to seek new employment or apply for DI benefits instead. Since in reality workers can endogenously quit jobs to obtain DI benefits, shifts in benefits supply might instead largely impact labor supply by inducing endogenous quits even absent adverse shocks. Below, we write a simple dynamic programming model to explore how elastic these

13. 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, visible in Figure I. It is estimated that approximately two-thirds of those terminated eventually requalified for benefits under a different impairment (cf. Lewin Group [1998]).

14. This set of facts does not imply that more recent disability beneficiaries are "shirking." As Bound and Waidmann stress [1992, 2000], disability is a continuous rather than a dichotomous medical state. A more expansive definition of disability will accommodate a greater range of illness. Bound and Waidmann [2000] provide evidence that the incidence of self-reported disability among males responds markedly to changes in the generosity of the disability program. Similarly, Baker, Stabile, and Deri [2001] find using matched medical records and health self-reports from Canada that individuals who are out of the labor force are more likely to report major medical ailments that are not reflected in objective health records. Notably, despite large increases in disability receipt, Burkhauser, Daly, Houtenville, and Nargis [2001] find that prevalence of self-reported disability by income decile has changed little over the past two decades.

responses—exit after shocks or direct quits—are likely to be to shifts in the supply of benefits.

We compare steady states of a discrete time Markov model in which individuals may be in one of three states in each period: employed; unemployed and seeking work; and seeking or receiving DI benefits (not participating). Employed workers receive per-period utility of employment $v(w,h)$, where the utility of work is increasing in wages and health: $v_w(\cdot)$, $v_h(\cdot) > 0$.[15]

Employed workers face a per-period hazard s of job loss and unemployed job seekers face a per-period hazard q of reemployment. As an alternative to employment, individuals may seek DI benefits. Applicants with health h qualify for benefits after one period with probability $p = p(h)$, and rejected applicants may reapply. Consistent with DI program rules, labor force participants are rejected with probability one.[16] Accepted applicants receive per-period income of d in perpetuity.

We assume a discount factor of $\beta < 1$ and set per-period utility of unemployment to zero.[17] To (further) simplify the analysis, we restrict attention to the case in which neither w nor h is time varying, and we consider a set of individuals who face a common probability of DI award conditional on application: $p(h_i) = p$ for all $_i$.

Using these parameters, the asset value of employment reduces (after some algebra) to the following expression:

$$(3) \quad V_E = \max\left[\frac{\beta p V_D}{1 - \beta(1 - p)}, \frac{(1 - \beta(1 - p)) \cdot v_i + \beta^2 p s V_D}{(1 - \beta(1 - s))(1 - \beta(1 - p))}, \right.$$
$$\left. \frac{(1 - \beta(1 - q)) \cdot v_i}{(1 - \beta)(1 - \beta(1 - q - s))}\right].$$

Each of the three terms inside of the brackets in (3) corresponds to one of three decision rules ("policies") that a worker may

15. 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], Burkhauser, Butler, and Weathers [1999], Kreider [1999], and Benitez-Silva et al. [2000] for theoretical and empirical models of the DI application decision.
16. Although it is possible that some DI applicants collect unemployment insurance benefits during the application process, this is technically illegal and likely rare since UI beneficiaries must demonstrate that they are active job seekers. UI recipients may of course seek DI benefits when UI is exhausted.
17. Setting per-period unemployment utility to zero is an assumption rather than a normalization. Provided that the flow utility of unemployment is less than the flow utility of receiving DI benefits, explicitly parameterizing the level of unemployment benefits adds complexity but does not change the core results.

pursue: apply for DI benefits immediately (I), apply conditional on job loss (C), or remain in the labor market after job loss (R).

The first bracketed term in (3) is the value of quitting employment directly to seek disability benefits. Note that if a worker selects this policy, she will find it optimal to quit employment immediately, to reapply for benefits if rejected, and to remain a beneficiary in perpetuity once accepted. The asset value of policy V_I is therefore the present discounted value of permanent benefits receipt $V_D = u(d)/(1 - \beta)$ discounted by the expected time from quit to award, where $u(d)$ is the per period utility of receiving DI benefits.

The second term is the value of remaining employed until exogenous job loss and then seeking disability benefits thereafter. We refer to this policy as "conditional application" and its value V_C incorporates the asset value of applying for and ultimately receiving DI benefits and the expected flow value of employment prior to DI application.

The third policy in brackets is the value of remaining in the labor market in perpetuity, V_R. Because workers pursuing this policy seek reemployment (not DI benefits) after job loss, the reemployment hazard q appears in this expression while the benefits award hazard p does not.
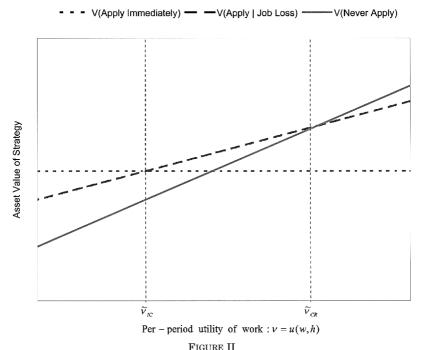
The three components of (3) are linked by the max[$\cdot$] operator because the value of employment for a given individual corresponds to the policy $\{I, C, R\}$ that yields the highest expected utility. For market participants arrayed along the distribution of v, the value of employment is the upper envelope of these three policies. Figure II depicts a simulation of the asset value of each policy as a function of v.[18] As is visible in the figure, these three slopes give rise to two thresholds labeled $\tilde{v}_{IC}$ and $\tilde{v}_{CR}$, in which the optimal policy shifts from (I) to (C) to (R) as a function of v. Solving for these thresholds yields

(4)

$$\tilde{v}_{IC} = \left(\frac{\beta p}{1 - \beta(1 - p)} \right) u(d), \ \tilde{v}_{CR} = \left(\frac{p u(d)}{q} \right) \left(\frac{1 - \beta(1 - q - s)}{1 - \beta(1 - p)} \right).$$

Equation (4) permits exploration of the question that motivates the model: how do changes in program parameters d and p

18. Parameter values used for the simulation are $\beta = 0.9$, $p = 0.5$, $u(d) = 1$, $s = 0.1$, and $q = 0.5$. Since p is assumed constant in the subpopulation depicted, variation in v can be viewed as arising from variation in w.

- - - V(Apply Immediately) ▬ ▬V(Apply | Job Loss) ▬▬▬V(Never Apply)

Per – period utility of work : $v = u(w,h)$

FIGURE II

The Choice of Disability Applicant Status—Immediately, Conditional, Never—as a Function of Earnings and Health

(the "supply" of benefits) and labor market parameters q and s (the "demand" for benefits) affect the share of workers pursuing each policy, in particular quitting work immediately to apply for benefits versus applying for benefits conditional on job loss?

Consider first the impact of demand conditions on labor force exit. An individual who is indifferent between immediate and conditional disability application is by definition indifferent to a change in the job loss hazard that hastens or retards the moment of application. Hence, $\tilde{v}_{IC}$ is independent of s. Because workers pursuing (I) and (C) do not reenter the labor market after exit, $\tilde{v}_{IC}$ is also independent of the reemployment hazard q. Consequently, changes in labor market conditions do not impact the size of the group that quits employment immediately to seek benefits (I). These parameters do, however, unambiguously impact the size of the conditional applicant group. Because higher s and lower q reduce the value of job search, more adverse labor

market conditions increase the share of workers applying for benefits in the event of job loss, policy (C).[19]

Next, consider an increase in the supply of benefits. Logically, increases in d and p shift both $\tilde{v}_{IC}$ and $\tilde{v}_{CR}$ rightward, raising the total share of workers who eventually seek benefits, both immediately and after job loss. The magnitude of the impact at the two thresholds differs, however:

$$(5) \quad \frac{\partial \tilde{v}_{CR}}{\partial d} > \frac{\partial \tilde{v}_{IC}}{\partial d} > 0 \text{ and } \frac{\partial \tilde{v}_{CR}}{\partial p} > \frac{\partial \tilde{v}_{IC}}{\partial p} > 0 \text{ and } \frac{\partial^2 \tilde{v}_{CR}}{\partial d \partial p} > \frac{\partial^2 \tilde{v}_{IC}}{\partial d \partial p} > 0.$$

Increases in d and p—and the interaction of the two—shift the conditional/remain-in-labor-force locus, $\tilde{v}_{CR}$, farther rightward than the immediate/conditional locus, $\tilde{v}_{IC}$. Provided that the density of v is weakly increasing between the previous $\tilde{v}_{IC}$ locus and the new $\tilde{v}_{CR}$ locus (as would be the case if the distribution of v were uniform), the size of the conditional applicant group will rise relative to the size of the immediate applicant group.

Finally, it is straightforward to show using the cross-partial derivatives of $\tilde{v}_{IC}$ and $\tilde{v}_{CR}$ that an increase in the supply of benefits interacts positively with adverse labor market conditions to increase the relative size of the conditional applicant group. 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 (and vice versa).[20]

On net, we find that "conditional" application is likely to be

19. It bears emphasis that this result—that direct quits are not influenced by labor market conditions—arises in part from our assumption that w is fixed; workers keep their current wage until job loss. While this assumption is clearly too strong, it is qualitatively consistent with well-known evidence that wages of incumbent workers are substantially more sheltered from labor market conditions than those of job seekers (cf. Beaudry and DiNardo [1991], Card and Hyslop [1997], and Kahn [1997]). We therefore expect the direct impact of s and q on w among employed workers to be second order. By contrast, it is quite likely that the expectation of w for *job losers* falls substantially at job displacement due to the destruction of specific capital or other incumbency related rents [Jacobson, LaLonde, and Sullivan 1993]. Broadening the model to permit expectations of w to depend on labor market conditions and the worker's employment/unemployment state would likely *increase* the differential attractiveness of disability application for job losers relative to employed.

20. More formally, the cross-partial derivatives $\partial^2 \tilde{v}_{CR}/\partial d \partial s$ and $\partial^2 \tilde{v}_{CR}/\partial p \partial s$ are strictly positive while $\partial^2 \tilde{v}_{CR}/\partial d \partial q$ and $\partial^2 \tilde{v}_{CR}/\partial p \partial q$ are strictly negative. All four corresponding cross-partial derivatives for the immediate/conditional threshold $\tilde{v}_{IC}$ are zero. Note that after sufficient time elapses, all "conditional applicants" exit the labor force, at which point the program exerts no further effect on their labor force participation. It is therefore appropriate to think of the model as applying to a single cohort of workers, with new cohorts entering the market in each period.

elastic to three forces: benefits supply, benefits demand, and the interaction of the two. By contrast, direct quits are only elastic to the first of these three forces and, in general, less elastic than is conditional application.[21] Our model therefore suggests that less stringent DI screening and higher replacement rates coupled with declining labor market prospects for the low skilled are likely to have increased the propensity of job losers to exit the labor force to seek disability benefits.

We explore two implications of the model below. In Section III we use the prediction from (5) that shifts in DI screening stringency should induce larger application responses where replacement rates are higher to test whether DI application rates and labor force participation among low skilled were differentially affected by the pre- and post-1984 shifts in DI screening in high versus low replacement rate states. In Section IV we explore the model's main implication: the responsiveness of DI application and labor force exit to adverse demand shocks should have risen secularly post DI liberalization.

III. DISABILITY BENEFITS AND LABOR FORCE PARTICIPATION: INSTRUMENTING FOR BENEFITS SUPPLY

We begin the empirical analysis by exploiting the disability retrenchment of 1979–1984 and subsequent liberalization of 1984–1998 to study the impact of shifts in the supply of disability benefits on labor force participation. A comparison of the two periods provides a useful contrast: per-capita DI receipt among nonelderly adults contracted at 0.10 percentage points annually during 1979–1984 and then expanded at 0.07 percentage points per year during the subsequent fourteen years. If the supply of disability benefits impacts labor supply, it should have had opposite impacts on labor force participation during these two time periods.

Suppressing subscripts, we write the conditional expectation of labor force participation as

$$(6) \quad E[LFP|X,w,d,p,h] = \alpha + \phi g(d,p) + \beta_1 w + \beta_2 h + X'\beta_3,$$

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21. A disadvantage of our focus on steady states is that we do not model the labor market impacts of unanticipated changes in parameter values that may induce immediate labor force exits that would not be visible in equilibrium. Consequently, our model should not be taken to imply that the impact of DI benefits on direct quits is negligible, only that the response of "conditional" applications is likely to be more elastic than direct exits.

where LFP is a dichotomous variable equal to one if an individual is a labor force participant, w is the opportunity wage, h is health, and X is a vector of individual characteristics. We denote the DI benefits "supply" faced by the potential labor force participant as $g(d,p)$, assumed to be increasing in both the replacement rate d and the odds of obtaining benefits conditional on application p. The coefficient of interest in (6) is ϕ, the impact of potential benefits on labor force participation.

There are a number of difficulties in estimating this equation with individual level data: we cannot typically observe both d and w for a given individual; objective measures of h are normally not available from survey data; and, as stressed by Bound [1989], because individuals with poor health typically experience declining wages and hence rising potential DI replacement rates, omission of h from (6) will bias estimates of ϕ.[22]

To surmount some of these biases, we estimate a state-level analog of (6) in first differences:

$$(7) \quad \Delta(LFP/Pop)_{j\tau} = \alpha + \phi\Delta DI_{j\tau} + \beta_1\Delta w_{j\tau} + \beta_2\Delta h_{j\tau} + \Delta X'_{j\tau}\beta_3 + \epsilon_{j\tau},$$

where j subscripts the 50 U. S. states excluding the District of Columbia and Δ_τ denotes the first difference operator over years t and τ. As a proxy for changes in the supply of disability benefits, we initially use the observed contemporaneous state-level changes in DI recipients per 1000 nonelderly population ages 25–64 ($\Delta DI_{j\tau}$). We subsequently instrument this measure. Since SSA does not report the educational attainment of DI beneficiaries, we use the total count of nonelderly recipients by state. OLS estimates of (7) are given in the first four columns of Table IV. Here, the dependent variable is the state-level annualized percentage point change in the labor force participation rate of the relevant demographic subgroup calculated from the merged monthly files of the Current Population Survey (CPS) for 1978 to 1998.[23]

Column 1 of Table IV indicates that states with more rapid

22. An additional difficulty is the possibility of unmeasured compositional shifts in the high school dropout population, which declined from 26 to 12 percent of overall population during 1979–1998. To (imperfectly) account for composition, we control for high school dropout age structure in each state in all models.
23. Models also control for the age distribution of the relevant demographic subgroups. We initially assume that, conditional on age and education, average wage and health changes are common across states and hence allow them to be absorbed by α. In Section VI we directly examine these (and other) alternative explanatory variables. A key assumption here is that health is not itself procyclical. Ruhm [2000] provides evidence that health is countercyclical.

TABLE IV

CHANGE IN DI ROLLS AND LABOR FORCE PARTICIPATION OF NONELDERLY ADULTS: OLS AND INSTRUMENTAL VARIABLES ESTIMATES

DEPENDENT VARIABLE: $100 \times$ ANNUALIZED CHANGE IN LABOR FORCE PARTICIPATION RATE

A. Δ Male labor force participation

	OLS estimates				IV estimates			
	High school dropouts		High school grad plus		High school dropouts		High school grad plus	
	78–84	84–98	78–84	84–98	78–84	84–98	78–84	84–98
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ DI Rolls/ 1000 Pop	−0.61	−0.61	−0.06	−0.07	−1.35	−0.51	−0.20	0.07
	(0.15)	(0.14)	(0.05)	(0.04)	(0.43)	(0.32)	(0.13)	(0.09)
Intercept	−1.24	0.25	−0.31	−0.09	−2.03	0.18	−0.47	−0.20
	(0.17)	(0.11)	(0.06)	(0.04)	(0.47)	(0.25)	(0.14)	(0.07)
R^2	0.43	0.59	0.19	0.13				
1st-stage Coefficient $\times 10^{-1}$					−0.77	0.51	−0.86	0.62
					(0.24)	(0.16)	(0.24)	(0.14)

B. Δ Female labor force participation

	OLS estimates				IV estimates			
	High school dropouts		High school grad plus		High school dropouts		High school grad plus	
	78–84	84–98	78–84	84–98	78–84	84–98	78–84	84–98
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ DI Rolls/ 1000 Pop	−0.22	−0.38	0.16	0.00	−1.01	−0.66	0.28	−0.14
	(0.18)	(0.15)	(0.10)	(0.08)	(0.47)	(0.31)	(0.25)	(0.15)
Intercept	−0.19	0.58	1.13	0.57	−1.04	0.59	1.27	0.65
	(0.22)	(0.13)	(0.11)	(0.07)	(0.52)	(0.14)	(0.27)	(0.13)
R^2	0.08	0.27	0.17	0.02				
1st-stage Coefficient $\times 10^{-1}$					−0.87	0.53	−0.75	0.60
					(0.25)	(0.14)	(0.25)	(0.12)

n = 50 U. S. states. Standard errors are in parentheses. Labor force participation rate is the fraction of nonelderly adults in the relevant gender/education category who are in the labor force (employed or unemployed). All estimates control for changes in the age distribution in the state population for the relevant gender/education group (ages 25–39 and 40–54 with 55–64 omitted). Age controls are demeaned in each time period. Estimates are weighted by mean state share of U. S. population in the two years used to form the dependent variable. Δ(DI Rolls/Pop) is annualized changes in DI recipients per 1000 state population ages 25–64. In OLS specifications, this variable is treated as exogenous. In IV specifications, it is instrumented by the potential DI earnings replacement rate for currently employed nonelderly adult males ages 25–61 at the seventy-fifth percentile of the state's earning replacement rate distribution in 1978. The potential replacement rate is calculated from the 1978 Current Population Survey/Social Security Earnings Records Exact Match file. Details on this calculation are provided in the Data Appendix.

declines in disability rolls per population during 1979 to 1984 experienced substantial contemporaneous (relative) *increases* in the labor force participation of high school dropout males. During this five-year interval, a one-tenth of a percentage point decline in disability receipt is associated with a rise in male high school dropout labor force participation of six-tenths of a percentage point.[24] Column 2 displays the analogous estimate for the labor force participation of high school dropout males during the post-liberalization era from 1984–1998. Although the variation in column 1 is driven by DI program contraction while variation in column 2 stems from program expansion, the estimated impact of DI benefits on labor force participation is near identical in each period.

Subsequent columns tabulate the relationship between DI receipt and labor force participation of males with at least a high school education. The point estimates are always small and typically insignificant. Although DI receipt increased among better educated adults after 1984 (Table II), this growth does not appear to have been accompanied by labor force exit.[25]

These OLS estimates will be biased if shifts in the supply of benefits due to program retrenchment and reform are correlated with state-level shifts in the demand for benefits (due, for example, to earnings and health shocks). To isolate the component of variation in DI rolls that is due to benefits *supply,* we exploit the progressivity of the DI benefits formula. This formula does not account for variation in regional wage levels and thus gives rise to substantial cross-state variation in potential replacement rates. As indicated by equation (5) of our model, shifts in screening stringency are predicted to have a greater impact on DI rolls where the level of the replacement rate is higher. Hence, the interaction between national shifts in the DI screening regime

24. Since high school dropouts comprise a far larger share of the disability recipient population than their share in the workforce, this coefficient is of reasonable magnitude: male high school dropouts averaged 7.9 percent of the labor force over 1984–1998 (CPS data) and comprised 25 percent of all disability recipients in 1999 (SIPP data).

25. Estimates found in Table 3 of the NBER Working Paper version of this paper [Autor and Duggan 2001] indicate that state-level changes in *SSI* receipt are unrelated to the labor force participation of high school dropouts (or those with higher education), in contrast to the results above for DI receipt. This pattern is consistent with the differing labor force histories of the beneficiaries of these two programs (see footnote 3).

and state replacement rates may serve as an instrument for the state-level shifts in the supply of DI benefits.[26]

Columns 5–8 of Table IV implement this instrumental variables approach. Using Social Security earnings records from the March 1978 CPS Exact Match file, we calculate that the median potential DI earnings replacement rate of currently employed males in 1978 ranged from a low of 18 percent in Alaska to a high of 38 percent in South Carolina, with a cross-state standard deviation of 2.6 percentage points. The first-stage coefficients tabulated at the bottom of columns 5–8 confirm that per-capita changes in DI rolls were substantially more negative in high replacement-rate states during 1978–1984 and, conversely, were substantially more positive in 1984–1998.[27]

The corresponding instrumental variables estimates of the impact of the supply of DI benefits on labor force participation by education group strongly corroborate the earlier OLS estimates. Inward shifts in the supply of DI benefits raised the labor force participation of low-skilled males and females during 1978–1984; outward supply shifts reduced them thereafter. IV estimates again show no impact of DI benefits on the labor force participation of more-skilled workers.

Since the growth in DI rolls was proportionately larger among females than males during 1984 to 1998, panel B of Table IV repeats these estimates for female labor force participation. As in panel A, we find that the supply of DI benefits reduces the labor force participation of high school dropout females and not those with higher education both during the retrenchment and reform periods. These results are especially striking in light of the fact that female labor force participation rose at all education levels over 1984–1998 (Table II).

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26. Provided of course that these interactions are not perversely correlated with state-level health and income shocks in *both* retrenchment and reform periods. This possibility seems particularly unlikely because the supply shifts are of opposite sign *for each state* in the retrenchment and reform eras. We provide further evidence on this point in Section VI.

27. The instrumental variable employed is the earnings replacement rate at the seventy-fifth percentile of the male potential DI replacement rate distribution calculated from the CPS-SSER. We use the earnings of males exclusively because of their high rate of labor force attachment, and use the seventy-fifth percentile replacement rate—typically corresponding to low earnings workers—because of our focus on labor supply of low skilled. Using the median replacement instead does not qualitatively impact our results. The interpretation of the first-stage coefficient in column 5 of panel A of the table is that for each additional percentage point of replacement rate in 1978, state disability rolls fell by 0.077 recipients per 1000 nonelderly adults per year during 1978–1984.

IV. Do Disability Benefits Affect How Low-Skilled Workers Respond to Adverse Employment Shocks?

The finding that shifts in DI benefits impact labor supply is consistent with two (not mutually exclusive) explanations: a once-and-for-all labor supply response, corresponding to an increase in direct quits in our model; or a secular shift in the sensitivity of labor force exit to adverse demand shocks, corresponding to an increase in "conditional exits." Our model predicts that the latter mechanism should have become increasingly significant following DI reform: reduced screening stringency, rising replacement rates, and declining demand for low-skilled workers should each have raised the relative attractiveness of DI application and labor force exit for workers facing adverse shocks.

To test this implication, we require a measure of plausibly exogenous labor demand shocks. Following the approach developed by Bartik [1991] and employed by Blanchard and Katz [1992] 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 $\hat{\eta}_{j\tau}$ for each state j between years t and τ as

$$(8) \qquad \hat{\eta}_{j\tau} = \sum_k \gamma_{jkt} \eta_{\not{j}k\tau},$$

where $\eta_{\not{j}k\tau}$ is the log change in two-digit industry k's employment share nationally and γ_{jkt} is the share of state employment in industry k in state j in the initial year. The subscript $\not{j}$ in $\eta_{\not{j}k\tau}$ indicates that each state's industry k employment is excluded in calculating the national employment share change.[28]

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. States that had a relatively large share of workers in declining industries will have predicted employment declines, while states that differentially employed

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28. In excluding own-state employment, our projected employment changes differ from the authors cited earlier. We found that including own-state employment substantially increased the predictive power of the employment projections, which raised a concern about a potential mechanical relationship. Also distinct from Bound and Holzer [2000], we compute $\eta_{\not{j}k\tau}$ as the change in log industry employment *shares* rather than levels. This makes the demand index strictly a relative measure.

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.

IV. A. Estimation Framework

We consider a modified form of equation (6) in which we write the conditional expectation of DI application for a displaced worker as

(9) $E[Apply|Job\ Loss,X,w,d,p,h]$

$$= \alpha + \sigma g(d,p) + \beta_1 w + \beta_2 h + X'\beta_3,$$

where $g(\cdot)$ again represents the disability supply function. Our model implies that $\sigma > 0$, greater benefits supply increases the propensity of adversely shocked workers to exit the labor force. As above, we estimate this equation using state-level data. Here, the move from individual to aggregate data is less straightforward. Ignoring demographic characteristics in (9) for the moment, we write the model for the flow of state DI applications in year t through τ as

(10) $(Apps/Pop)_{j\tau} = \alpha + \lambda_t \Delta(Emp/Pop)_{j\tau} + \epsilon_{j\tau},$

where λ_t measures the direct impact of employment losses on DI application flows in year t. Similarly, we model labor force exit as

(11)

$\Delta(NILF/Pop)_{j\tau} = \gamma + \theta_t \Delta(Emp/Pop)_{j\tau} + \lambda_t \Delta(Emp/Pop)_{j\tau} + \upsilon_{j\tau}.$

In this equation declines in labor force participation come from two sources: workers who withdraw from the labor force but do not seek DI benefits (measured by θ_t); and job losers who leave the labor force to seek DI benefits (λ_t). As the equation underscores, our data do not permit a direct estimate of σ in (9). Instead, estimates of (11) will yield a reduced form that combines both sources of labor force exit:

(12)

$\Delta(NILF/Pop)_{j\tau} = \alpha_0 + \pi_t \Delta(Emp/Pop)_{j\tau} + \upsilon_t,$ where $\pi_t = \lambda_t + \theta_t.$

Given this constraint, we estimate equations (10) and (12) separately in both the pre- and post-DI liberalization eras to

obtain estimates of $\Delta\lambda_\tau$ and $\Delta\pi_\tau$, the change in the responsiveness of DI applications and labor force exit to demand shocks of given magnitude. With these estimates, we may distinguish three cases. First, if DI application sensitivity has risen but labor force exit propensity has not ($\Delta\hat{\lambda}_\tau > 0$ and $\Delta\hat{\pi}_\tau \approx 0$), this would suggest that rising DI benefits supply has increased program take-up but not impacted labor force exit. Conversely, if $\Delta\hat{\lambda}_\tau \approx 0$ and $\Delta\hat{\pi}_\tau > 0$, this would indicate that labor force exit has risen for given shocks but that this change is unrelated to the supply of DI benefits. Finally, if $\Delta\hat{\lambda}_\tau \approx \Delta\hat{\pi}_\tau > 0$, application and labor force exit propensity have risen conformably, we will view this as evidence that the rising supply of DI benefits has increased the propensity of adversely shocked workers to withdraw from the labor force to apply for DI benefits. A limitation of this approach is that it does not separately identify the impacts of rising replacement rates and declining screening stringency on labor force exit propensity. The effect that we identify will be a sum of these structural components.

IV. B. *Estimation: Adverse Employment Shocks and DI Application Flows*

Estimates of equation (10), the impact of employment shocks on DI application flows, are given in Table V. The dependent variable in this table is the annualized state level flow of unique DI applications per capita during 1978–1984 and 1984–1998. The independent variable is the annualized projected state log employment shock measure $\hat{\eta}_{j\tau}$ for the corresponding time interval. Models also include time dummies and control for changes in the age and gender distribution of nonelderly adult state residents. As above, we initially make the assumption that wage and health variables are uncorrelated with employment shocks but examine their impact directly in Section VI.

Column (1) of Table V provides an estimate of equation (10) for the pre-DI liberalization period of 1978–1984. The point estimate of -0.13 indicates that a 1-percent projected log state employment contraction yielded 1.3 additional DI applications per 1000 nonelderly adults over these five years. The subsequent column pools data for the entire time period, adding an interaction between the shock measure, $\hat{\eta}_{j\tau}$, and a post-1984 dummy variable. The coefficient on the interaction term implies that the impact of adverse shocks on DI application flows grew threefold from the retrenchment to the reform era. Adding state dummies

TABLE V

IMPACT OF EMPLOYMENT LOSSES ON DI APPLICATIONS FLOWS 1978–1998: REDUCED-FORM AND INSTRUMENTAL VARIABLES ESTIMATES

DEPENDENT VARIABLE: ANNUALIZED FLOW OF DISABILITY APPLICANTS PER NONELDERLY ADULT

A. OLS reduced-form estimates: impact of predicted Δ(Emp/Pop) on DI Adds/Pop

	Long changes			Stacked 3-yr diffs		
	(1)	(2)	(3)	(4)	(5)	(6)
	78–84	78–98	78–98	78–84	78–98	78–98
Δ Emp/Pop	−0.13 (0.06)	−0.13 (0.07)	−0.11 (0.05)	−0.12 (0.04)	−0.11 (0.04)	0.00 (0.02)
Δ Emp/Pop × Post-1984		−0.34 (0.09)	−0.26 (0.06)		−0.17 (0.05)	−0.05 (0.02)
State dummies	No	No	Yes	No	No	Yes
R^2	0.29	0.32	0.97	0.31	0.35	0.90
n	50	100	100	100	350	350

B. IV Estimates: impact of high school dropout Δ(Emp/Pop), instrumented by predicted Δ(Emp/Pop), on DI Apps/Pop

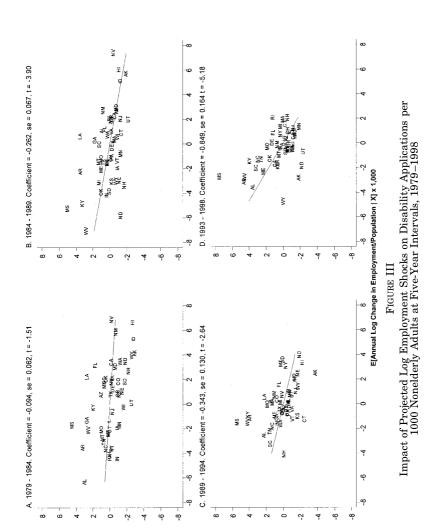|	Long changes			Stacked 3-yr diffs		
	(1)	(2)	(3)	(4)	(5)	(6)
	78–84	78–98	78–98	78–84	78–98	78–98
Δ Emp/Pop	−0.25 (0.20)	−0.30 (0.22)	−0.05 (0.13)	−0.14 (0.07)	−0.12 (0.07)	0.01 (0.02)
Δ Emp/Pop × Post-1984		−0.68 (0.67)	−0.34 (0.23)		−0.15 (0.11)	−0.05 (0.03)
State dummies	No	No	Yes	No	No	Yes
1st-stage coef. (main effect)	0.51 (0.25)	0.52 (0.29)	1.07 (0.69)	0.85 (0.24)	0.90 (0.26)	1.26 (0.33)
1-stage coef. (interaction)		0.54 (0.28)	1.69 (1.27)		1.01 (0.30)	1.62 (0.42)
F-statistic	3.97	7.87	1.19	12.02	17.26	11.41
P-value	0.05	0.00	0.31	0.00	0.00	0.00
n	50	100	100	100	350	350

Standard errors in parentheses account for clustering of state level observations in models that contain repeated observations per state and exclude state dummies (specifications 2, 4, 5). Dependent variable is the annualized flow of unique DI applications per nonelderly state population (ages 25–64) over relevant time interval: six years in specification 1; six and fourteen years in specifications 2 and 3; and 3-year intervals in specifications 4–6. Data points for stacked 3-year changes correspond to years 78–81, 81–84, 84–87, 87–90, 90–93, 93–96, and 95–98. In regressions covering the years 78–98 (specifications 2, 3, 5, and 6), observations for 78–84 and 84–98 are stacked. The Δ(Emp/Pop) variable in Panel A (and its interaction with the post-1984 dummy) is the predicted state level change in the employment to population rate, calculated as the national level change in two-digit industry log employment shares (excluding own-state employment) over the relevant interval projected onto initial state industry shares. See equation (8) of the text for details of this calculation. The Δ(Emp/Pop) variable used in Panel B is the observed change in the employment to population rate of high school dropouts instrumented by the predicted employment change measure in Panel A (and its interaction with a post-1984 dummy). All models include an intercept, time dummies, and controls for annualized state level changes in population shares by age and gender: male/female × ages 25–39/40–54/55–65 (six categories, one omitted). Predicted employment change measure, employment to population ratio of high school dropouts, and state age × gender distribution variables are calculated from merged CPS monthly files for years 1979–1998. Disability application counts are provided by the Social Security Administration.

to control for fixed state differences in DI application flows only slightly reduces the magnitude of the estimated growth, which remains highly significant.

To potentially improve upon the precision of these estimates, we reestimate these models using stacked data on application flows and projected demand shocks by state at three-year intervals. These models found in the subsequent three columns provide qualitatively similar results. The estimated growth in DI application sensitivity is not as sizable in the stacked models, but remains quite large and significant.[29] To illustrate the source of this result, Figure III plots the conditional relationship between estimated log demand shocks (multiplied by 10^3) and DI applications per 1000 population for four five-year subperiods of our sample. Consistent with the contemporaneous rise in replacement rates and the decline in screening stringency post 1984, the slope of the demand shock-DI application locus increases secularly in each adjacent panel. By 1993–1998 we estimate that this slope has increased sevenfold.

We have performed numerous robustness tests to verify these relationships. If the disability application variable is replaced with the flow of DI awards or the change in DI recipients per population, we find qualitatively similar patterns [Autor and Duggan 2001]. We have also utilized county level DI receipt data to explore the sensitivity of disability rolls over 1984–1998 to *within-state* variation in local demand shocks generated by variation in county-level industrial structure. Our estimates confirm that within states, counties experiencing relatively negative predicted employment shocks saw substantially faster growth in DI rolls over 1984–1998.

Panel B of Table V presents a variant of the models above in which we estimate the impact of changes in the overall high school dropout employment to population ratio on DI application flows while instrumenting high school dropout employment with the projected demand shock measure. Because these models only allow demand shocks to impact DI applications through their covariance with the high school dropout employment to popula-

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29. The stacked estimates are not strictly preferable since they may accentuate the importance of measurement error in the demand instrument at short intervals, particularly in models with state fixed effects. To implement these estimates, we use changes for 1978–1981 and 1981–1984 for the 1978–1984 period. For the 1984–1998 estimates, we use changes for 1984–1987, 1987–1990, 1990–1993, 1993–1996, and (due to sample truncation) 1995–1998.

FIGURE III

Impact of Projected Log Employment Shocks on Disability Applications per
1000 Nonelderly Adults at Five-Year Intervals, 1979–1998

tion ratio, it is notable that IV estimated impacts of demand shocks on DI applications are closely comparable to the stacked OLS models in panel A. High school dropout employment levels and DI application flows appear to respond to the same component of the demand shock measure.[30]

IV. C. *Estimation: Adverse Employment Shocks and Labor Force Exit of High School Dropouts*

Since essentially all disability applicants are labor force nonparticipants, the finding of rising DI application sensitivity for given shocks has two potential explanations. One is that since 1984, employment shocks have increasingly spurred labor force nonparticipants to apply for disability benefits, perhaps in anticipation of difficulty finding work at labor market reentry. A second is that the share of job losers exiting the labor force to seek disability benefits has risen. In terms of equation (12), these cases correspond to $\Delta\pi_\tau \approx 0$ versus $\Delta\pi_\tau > 0$.

We distinguish these possibilities by estimating (12) using projected demand shocks to instrument for net employment losses at the state level. Using the identity that $\Delta(Emp/Pop)_{j\tau} = -\Delta(Unemp/Pop)_{j\tau} - \Delta(NILF/Pop)_{j\tau}$, we can decompose the impact of employment shocks into the component accruing to unemployment and the component accruing to labor force exit. To facilitate comparison to the DI application regressions, we present results for both male and female high school dropouts combined and disaggregated estimates by gender. These are found in Table VI.

During 1978–1984 a one-percentage-point shock to the employment to population ratio of high school dropouts induced a 0.46 percentage point increase in nonparticipation per population and, by implication, a 0.54 percentage point increase in unemployment per population. In the succeeding fourteen years, shocks to employment of equal magnitude led to substantially larger declines in labor force participation. In both 1984–1990 and 1990–1998, we find that essentially all declines in high

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30. Because the first stage of the long change models in columns (1)–(4) is weak, we focus discussion on the stacked models. A significant limitation of these models is that our DI application measure is not disaggregated by education level, and hence it is unlikely that the data satisfy the implied exclusion restriction that demand shocks impact DI applications only through high school dropout employment changes.

TABLE VI

INSTRUMENTAL VARIABLES ESTIMATES OF THE IMPACT OF EMPLOYMENT LOSSES ON LABOR FORCE EXIT OF HIGH SCHOOL DROPOUTS, 1979–1998
DEPENDENT VARIABLE: CHANGE IN HIGH SCHOOL DROPOUT NONPARTICIPATION RATE

	A. Males and females combined							
	1978–1984		1984–1990	1990–1998	1984–1998		1978–1998	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ Emp/Pop	−0.46	−0.73	−1.47	−0.93	−1.10	−0.90	−0.57	−0.58
	(0.23)	(0.16)	(0.52)	(0.21)	(0.21)	(0.18)	(0.19)	(0.19)
Δ Emp/Pop × (85–98 dummy)							−0.34	−0.36
							(0.16)	(0.19)
State dummies	No	Yes	No	No	No	Yes	No	Yes
n	100	100	100	150	250	250	350	350

	B. Estimates by gender							
	1978–1984		1984–1998		1978–1998		1978–1998	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Males	Females	Males	Females	Males	Females	Males	Females
Δ Emp/Pop	−0.45	−0.44	−0.98	−1.23	−0.54	−0.60	−0.54	−0.49
	(0.21)	(0.40)	(0.19)	(0.34)	(0.16)	(0.29)	(0.22)	(0.22)
Δ Emp/Pop × (85–98 dummy)					−0.43	−0.60	−0.49	−0.56
					(0.25)	(0.33)	(0.27)	(0.36)
State dummies	No	No	No	No	No	No	Yes	Yes
n	100	100	250	250	350	350	350	350

Standard errors in parentheses account for clustering of state level observations in models excluding state dummies. All data points are stacked three-year state level changes in the relevant measure for years 78–81, 81–84, 84–87, 87–90, 90–93, 93–96, and 95–98. Models also include year dummies and control for changes in the age distribution of state high school dropout population (ages 25–39 and 40–54, with 55–64 omitted) for the relevant gender group. Estimates are weighted by state share of U. S. population in each year. Instrumental variable used for change in employment to population ratio of high school dropouts in all models is projected state level log employment shock for relevant time interval and, in pooled 1978–1998 specifications, its interaction with a post-1984 dummy. See equation (8) in the text for further details on projected employment shock measure.

school dropout employment led to increases in nonemployment rather than unemployment.

To control for persistent state trends in labor force exit that would bias the results if correlated with the demand shock measure, we add state fixed effects to models in columns (2), (6), and (8). Their addition has little impact on the pattern of estimates. To further explore robustness, columns (7) and (8) pool data for 1978–1998, adding an interaction between the employment loss measure and a post-1984 dummy variable. The pooled estimates confirm a significant increase in the labor force exit propensity of

displaced high school dropouts after 1984, on the order of 35 percent. Inclusion of state fixed effects raises the standard errors slightly, but does not alter the findings. As shown in panel B, the pattern of results is essentially identical for both genders.

It bears emphasis that these estimates do not imply that commencing in 1984, all—or even most—high school dropouts who lost jobs exited the labor force. Since our demand shock captures net rather than gross employment declines over comparatively long three-year intervals, displaced high school dropouts who obtain reemployment within the measurement window are not detected. In addition, the shock measure does not impact the labor market exclusively through job loss. Adverse shocks will induce some subset of workers to choose to exit employment and, perhaps more importantly, slow the prevailing rate of reentry of nonparticipants into the labor force. Similarly, some share of those exiting may be previously unemployed workers whose re-employment prospects were harmed by entry of new unemployed. Each of these forces will reduce the employment to population ratio and yield corresponding increases in nonparticipation. What is unambiguous from Table VI is that unemployment rates of low-skilled workers have become significantly less sensitive to adverse demand shocks while their rate of labor force nonpartici-pation has become substantially more so.

Could the increased generosity of the DI program plausibly account for this behavioral change? To make this calculation, note that $\Delta\hat{\lambda}_\tau$ in Table V measures the impact of employment shocks on DI applications for the entire nonelderly population, while $\Delta\hat{\pi}_\tau$ in Table VI is measured only for high school dropouts. To make the magnitudes commensurate, we calculate

$$(13) \qquad \Delta\hat{\lambda}_\tau^D = (1/\delta)\xi\Delta\hat{\lambda}_\tau,$$

where δ is the high school dropout share of population in the years of estimation, equal to 16 percentage points in our sample, and ξ is the share of the total flow of DI applications submitted by high school dropouts, which we estimate to be 47 percent.[31] Our estimate of $\Delta\hat{\lambda}_\tau$ from Table V ranges from 0.05 to 0.17, i.e., the fraction of all nonelderly adults applying for DI benefits in response to a one log

31. The rate of DI receipt among high school dropouts averaged 4.6 times that of those with at least a high school degree during 1984–1999 (Table II). Assuming that this ratio carried over to applications, high school dropouts would have accounted for 47 percent of all DI applications received in this period.

point adverse demand shock rose by 0.5 to 1.7 applicants per 1000 adults post liberalization. In this equation, $1/\delta$ rescales the estimated increase in DI applications per population to an increase per *high school dropout* population, while the parameter ξ adjusts this estimate downward to account for the fact that only part of the growth in DI applications is due to high school dropouts.

Substituting into (13), we obtain that the implied increase in high school dropout labor force exit propensity for a one log point employment contraction, $\Delta\hat{\lambda}_\tau^D$, ranges from 0.15 to 0.50. That is, if the increase in high school dropout application propensity were entirely accounted for by labor force leavers, the estimated increase in application sensitivity implies a 15 to 50 percentage point increase in their conditional probability of labor force exit. This estimate is in close accord with the findings in Table VI. Specifically, columns (7) and (8) of Table VI indicate that the share of high school dropouts exiting the labor force in response to a one unit shock post-liberalization increased by 34 percentage points ($\Delta\hat{\pi}_\tau^D = 0.34$). Notably, the estimates of $\Delta\hat{\lambda}_\tau^D \in [0.15, 0.50]$ from (13) fall symmetrically at one-standard error above and below this estimate. Based upon this (necessarily) rough calculation, we conclude that $\Delta\hat{\lambda}_\tau \approx \Delta\hat{\pi}_\tau$: the increasing supply of DI benefits post-1984 can plausibly account for both greater DI application flows for given shocks and greater propensity of adversely shocked low-skilled workers to exit the labor force. Future analyses using micro-data on the employment and DI application behavior of low-skilled, displaced, and discouraged workers should be able to improve on the precision of this estimate.

V. IMPLICATIONS FOR THE AGGREGATE LABOR MARKET

The reduced sensitivity of high school dropout unemployment to adverse labor demand shocks implies that the share of high school dropouts unemployed at present is lower than it would have been if the supply of DI benefits had not shifted outward after 1984. Consider that the typical nonelderly (that is, age 25–64) high school dropout in 1984 was employed in an industry that over the next fourteen years declined by 8.53 percentage points as a share of aggregate employment. Combining this negative relative demand contraction with the estimate of $\Delta\hat{\pi}_\tau^D = 0.34$, we calculate that the share of nonelderly adult high school dropouts who were unemployed would have been 2.9 percentage points higher in 1998 ($8.53 \times 0.34 = 2.90$) but for the
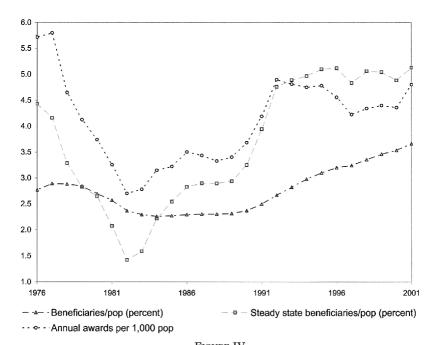
liberalization of DI that occurred in 1984. Since nonelderly high school dropouts accounted for 12.1 percent of the nonelderly adult population in 1998, this suggests that the aggregate unemployment rate would have been approximately half (0.44) of a percentage point higher in 1998 if it were not for DI liberalization.[32] Given that the overall unemployment rate of nonelderly adults in this year was 3.4 percent, this appears a sizable deviation from the conventional metric.

Although we have focused on the 1984–1998 period, it would be inaccurate to conclude that the labor market impact of the DI program had culminated by 1998. As is shown in Figure IV, though per-capita DI awards peaked in 1992 and declined after 1995 (that is, until the 2000 recession), the rate of DI receipt grew steadily throughout the late 1990s. These facts—rising recipients despite declining awards—indicate that the DI recipient population was below steady state size. Due to post-1984 declines in DI beneficiary mortality and retirement rates (Figure I), the expected duration of DI receipt for new entrants considerably exceeded that of previous cohorts, implying an increase in steady state program population.

To make a simple projection of this long-run size, we use the fact that in steady state, benefit terminations must equal benefit awards. Hence, $P^*\lambda = A$, where P^* is the steady state population, λ is the termination rate, and A is the number of new awards. Drawing on data from SSA's *Annual Statistical Supplement* [various years], Figure IV depicts estimates of P^* normalized by population for the years 1976 to 2001. Following the 1984 program liberalization, the steady state program size rapidly doubled from its 1982 low. At the onset of the 1991 recession, award rates increased by another 50 percent and the projected program size rose accordingly. By 1995 the projected steady state program size had reached 5.1 percent of nonelderly adults, where it largely remained for the balance of the decade.[33]

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32. Because the unemployment rate is denominated by labor force and not by population, the estimated impact using the overall nonelderly adult (ages 25–64) participation rate of 80.3 percent is $(2.90 \times 0.121)/0.803 = 0.44$.

33. By using average termination rates, the calculation does not adjust for the fact that more recent DI entrants are likely to have lower mortality and retirement rates than incumbent recipients. Accounting for these cohort effects would likely raise the steady state projection considerably. The projected steady state series in Figure VII is imputed in 1997 using the average of 1996 and 1998 to discount the one-time termination of beneficiaries whose impairment was Drug and Alcohol Addiction (see Figure I and footnote 13).

FIGURE IV
DI Awards and Current and Steady State Recipients per Nonelderly Adult
Population, 1976–2001
*Source for beneficiaries and awards: Social Security Bulletin: Annual Statistical
Supplement* [various years] denominated by estimates of adult population ages
25–64 from Current Population Survey. Projected steady state beneficiaries per
population is annual awards per adult population ages 25–64 divided by annual
termination rate.

Given that 3.7 percent of nonelderly adults received DI
benefits in 2001, this calculation augurs another 40 percent
growth in DI program size relative to population over approxi-
mately the next decade. Notably, application rates per popu-
lation rose an additional 22 percent between 1999 and 2001,
indicating the potential for another rise in steady state pro-
gram size. The DI recipient population may never reach this
projected size, however. As the figure indicates, steady state
program size in 1976 *exceeded* its 2001 steady state level. It
was SSA's recognition of this fact that galvanized the DI pro-
gram retrenchment and subsequent liberalization that form
the backdrop to our study.

VI. ALTERNATIVE EXPLANATIONS: DECLINING WAGES AND HEALTH,
RISING IMMIGRATION, INCARCERATION, AND UNEMPLOYMENT BENEFITS

Our conclusions above must be tempered to the extent that post-1984 increases in the DI application and labor force exit propensities of low-skilled workers are explained by factors other than shifts in the supply of DI benefits. Plausible alternative explanations for the cross-state patterns of labor force decline that are the focus of our study include real wage declines, differential health trends, rising immigration and incarceration rates, and shifts in state Unemployment Insurance (UI) benefits. We explore each of these alternatives below. Means and cross-state standard deviations of these measures are given in the first column of Table VII.

We begin with wages. As argued by Juhn [1992] and Juhn, Murphy, and Topel [1991], a sizable elasticity of labor force participation coupled with declining real wages could explain the substantial decline in labor force participation of low skilled males in the 1980s and 1990s. To be clear, we view falling wages as a complement rather than a substitute for the role of the DI program in reducing labor force participation. Indeed, an important component of the rise in replacement rates for low-skilled workers is the interaction between declining wages and the indexation and progressivity of the DI benefits formula. For declining 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 had larger declines in labor force participation.

To explore this hypothesis, we use CPS Merged Outgoing Rotation Group earnings data for 1984 and 1998 to construct two measures of the change in the opportunity wage faced by low-skilled workers in each state: the change in the real fixed-weighted median log wage of high school dropout males ages 25–64 and the change in the real fixed-weighted log wage of prime age males 25–54 at the twenty-fifth percentile.[34] Because the declining wage hypothesis implies that workers with the lowest potential earnings are least likely to work, we follow

34. In an effort to maximize the explanatory power of alternative hypotheses, we chose the more inclusive measure of wages for male high school dropouts (i.e., ages 25–64 rather than 25–54). Empirically, the wage measure for the younger group of high school dropouts works similarly but has less predictive power in labor force participation regressions.

TABLE VII
DETERMINANTS OF GROWTH IN DI ROLLS, 1984–1998
DEPENDENT VARIABLE: $100 \times$ ANNUALIZED CHANGE IN SHARE OF NONELDERLY ADULTS RECEIVING DI

	Means	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
DI replacement rate	0.377 (0.028)	0.67 (0.15)		0.46 (0.17)	0.48 (0.22)	0.71 (0.15)	0.49 (0.15)	0.53 (0.16)	0.63 (0.16)	0.54 (0.14)	0.68 (0.15)	0.65 (0.15)	0.50 (0.16)	0.46 (0.17)
E[ΔEmp/Pop] × 10	−0.017 (0.017)		−1.01 (0.24)	−0.66 (0.26)	−1.20 (3.07)								−0.65 (0.25)	−0.41 (0.34)
E[ΔEmp/Pop] × rep. rate	−0.006 (0.007)				0.14 (0.78)									
Δ HS dropout median wage	−0.186 (0.120)					−0.06 (0.03)							−0.06 (0.03)	−0.04 (0.03)
Δ 25th pentile wage	0.084 (0.121)						0.10 (0.03)							
Mortality rate '84 × 100	0.510 (0.063)							0.15 (0.08)						0.08 (0.08)
Δ Mortality rate × 100	−0.123 (0.038)								0.10 (0.12)					
Δ Share foreign born	0.049 (0.038)									−0.36 (0.09)				−0.14 (0.16)
Δ Incarceration rate × 10	0.046 (0.020)										−0.32 (0.18)			−0.14 (0.16)
Δ UI replacement rate	−0.025 (0.118)											0.03 (0.03)		
Intercept	0.078 (0.032)	−0.17 (0.06)	0.10 (0.02)	−0.08 (0.07)	−0.09 (0.08)	−0.19 (0.06)	−0.11 (0.06)	−0.21 (0.06)	−0.15 (0.07)	−0.09 (0.06)	−0.16 (0.06)	−0.16 (0.06)	−0.10 (0.07)	−0.11 (0.07)
R^2		0.45	0.44	0.53	0.53	0.50	0.56	0.50	0.46	0.59	0.49	0.48	0.57	0.63
N		50	50	50	50	50	50	50	50	50	50	49	50	50

Standard errors are in parentheses (cross-state standard deviations given for column of means). All models control for changes in age by gender shares of state population ages 25–64 (males/females by ages 25–39/40–54/55–64) and are weighted by states' average share of population in beginning and end year.

DI Replacement rate is the seventy-fifth percentile DI potential earnings replacement rate of employed nonelderly males calculated from the 1978 Current Population Survey–Social Security Earnings Records Exact Match file.

E[ΔEmp/Pop] is the annualized state projected log employment change for 1984–1998. See equation (8) in the text.

ΔMale high school dropout median wage is the change in the fixed weighted median wage of all male high school dropout state residents ages 25–64 excluding the self-employed, those working without pay, and employed workers for whom a wage could not be calculated. Weights correspond to 1984 state male population shares for ages 25–64 in ten year brackets (four categories).

Δ twenty-fifth percentile wage is the analogous fix weighted wage for male prime-age residents ages 25–54.

Mortality rate is the number of nonaccidental, nonhomicidal deaths among ages 25–64 divided by the average of state population 25–64 in the previous and current year. Δ Mortality rate is the change in this number.

Δ Share foreign born is the 1984 to 1998 change in share of population 25–64 that is foreign born calculated from the 1998 Current Population Survey and linearly interpolated for the year 1984 using the 1980 and 1990 Census PUMS 1 percent samples.

Δ Incarceration rate is the change in state prison population (excluding jails) per capita 18–64 from the Bureau of Justice Statistics.

Δ UI Replacement Rate is the 1984–1999 change in state UI replacement rate for a full-time minimum wage worker from the *Green Book: Overview of Entitlement Programs* [U. S. House of Representatives, various years] and is not available for Michigan.

Chandra [2002] in including nonparticipants in the quantile cal-
culations, thus assuming that the potential earnings of nonpar-
ticipants fall below the median (or twenty-fifth) percentile of the
conditional state wage distribution. Median potential wages of
male high school dropouts declined 18.6 log points between 1984
and 1998. In the same period there was a modest rise at the
twenty-fifth percentile for prime age male workers, reflecting
strong earnings growth in the late 1990s [Katz and Krueger
1999].

Although aggregate health improved during the two decades
of our sample, differential cross-state health trends could be in
part responsible for our findings. To test this possibility, we
exploit an objective health measure: the mortality rate (excluding
accidental deaths and homicides) among state residents ages
25–64.[35] We control for both levels and changes of this measure
since each should impact the response of state DI rolls to a
reduction in DI screening stringency. Mortality rates declined by
one-quarter between 1984 and 1998.

A third potential confound is immigration. Because less-
educated immigrants typically have greater labor force attach-
ment than native-born high school dropouts [Trejo 2001] and are
not eligible for disability benefits until five years after arrival,
geographically concentrated immigrant inflows during the 1980s
and 1990s could induce negative cross-state correlations between
high school dropout labor force participation and DI receipt. We
measure nonelderly foreign-born population using data from the
1980 and 1990 Census of Populations and the 1998 CPS. The
foreign-born share of nonelderly adults rose by close to 5 percent-
age points during 1984–1998, with sizable cross-state variation.

A fourth concern is growing incarceration. Because the in-
carcerated population is included in neither the numerator nor
denominator of our labor force calculations, it is not intrinsically
a source of bias. However, potential criminals have below average
rates of labor force participation relative to other high school
dropouts [Western, Kling, and Weiman 2001]. If states with low
DI growth also experienced differential increases in incarceration
over 1984–1998, this would bias our estimates. We assemble
Bureau of Justice Statistics data on state prison populations

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35. We focus on an objective measure due to the evidence in Bound and
Waidmann [1992, 2000] that male health self-reports closely track expansions and
contractions in the supply of disability benefits.

(excluding local jails), denominated by state population ages 18–64 [U. S. Department of Justice 2002]. Consistent with other authors [Freeman 1991; Katz and Krueger 1999; Western and Pettit 2000], we find that incarceration rates rose by a sizable 0.46 percentage point during 1984–1998.

Finally, the UI system provides an alternative (albeit short-term) source of wage replacement for displaced workers. Declines in the value of UI benefits during the 1980s and 1990s [Anderson and Meyer 1997] could potentially induce displaced workers to seek DI instead. We measure UI generosity as the 1984–1999 change in the state UI replacement rates for a full-time minimum wage worker using data from the U. S. House of Representatives *Green Book* [various years]. Although the mean UI replacement rate did not change appreciably between 1984 and 1998, some states made substantial adjustments.

We present two sets of estimates with these variables. Table VII shows the relationship between each measure and the annualized percentage point growth in DI rolls per nonelderly state population over 1984–1998. Table VIII presents analogous estimates for the annualized change in high school dropout labor force participation in the same period. Because results are for the most part comparable for the two dependent variables, we discuss them in combination.

The first three columns in each table include our two main explanatory variables: DI replacement rates (as a proxy for shifts in DI benefits supply) and projected employment shocks. Each has a significant impact on DI growth and high school dropout labor force nonparticipation. When entered in combination, the projected shock measure is not precisely estimated in the labor force participation equation while the replacement rate measure remains significant at the 10 percent level. Since our model suggests that shocks should have larger impacts on program uptake and labor force participation where replacement rates are higher, we interact the replacement rate and shock measure in column (4). With only 50 data points, the interaction term is very imprecisely estimated.

Subsequent columns of Tables VII and VIII explore the relationship between the two outcome measures and the other candidate variables, each entered in combination with the replacement rate. Higher initial mortality (though not its change), declines in median wages of high school dropouts, and decreases in the share of foreign born, each contributed to growth in state

TABLE VIII
DETERMINANTS OF CHANGE IN HIGH SCHOOL DROPOUT LABOR FORCE PARTICIPATION, 1984–1998
DEPENDENT VARIABLE: 10 × ANNUALIZED CHANGE IN HIGH SCHOOL DROPOUT LABOR FORCE PARTICIPATION RATE

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
DI replacement rate	−0.35 (0.15)		−0.26 (0.15)	−0.21 (0.20)	−0.33 (0.14)	−0.28 (0.15)	−0.32 (0.16)	−0.34 (0.15)	−0.30 (0.14)	−0.32 (0.15)	−0.38 (0.16)	−0.26 (0.15)	−0.34 (0.18)
E[ΔEmp/Popl] × 10		0.55 (0.22)	0.31 (0.23)	−0.36 (2.68)								0.31 (0.23)	−0.15 (0.36)
E[ΔEmp/Popl] × rep. rate				0.20 (0.69)									
Δ HS dropout median wage					0.08 (0.03)							0.07 (0.03)	0.04 (0.03)
Δ 25th pcntile male wage						−0.07 (0.03)							
Mortality rate '84 × 100							−0.03 (0.07)						0.00 (0.06)
Δ Mortality rate × 100								−0.04 (0.10)					
Δ Share foreign born									0.31 (0.09)				0.30 (0.17)
Δ Incarceration rate × 10										0.26 (0.19)			0.00 (0.20)
Δ UI replacement rate											0.02 (0.03)		
Intercept	0.12 (0.06)	0.00 (0.01)	0.10 (0.06)	0.08 (0.08)	0.13 (0.06)	0.10 (0.06)	0.13 (0.06)	0.11 (0.06)	0.10 (0.05)	0.10 (0.06)	0.13 (0.06)	0.10 (0.06)	0.11 (0.06)
R^2	0.39	0.39	0.42	0.43	0.47	0.44	0.39	0.39	0.51	0.41	0.40	0.49	0.54
—	50	50	50	50	50	50	50	50	50	50	49	50	50

Standard errors are in parentheses. All models control for changes in age distribution of stage high school dropout population ages (25–39/40–54/55–64 with one category omitted) and are weighted by states' average share of population in beginning and end year. See notes to Table VIII for variable definitions.

DI rolls and declines in high school dropout labor force participation during these fourteen years.[36] We do not find a significant relationship between incarceration or UI replacement rates and either DI receipt or high school dropout labor force participation. Notably, none of these measures greatly affects the estimated impact of the replacement rate on growth in DI rolls and declines in labor force participation. The final specifications in each table combine groups of candidate variables shown to be significant in previous columns. In all cases, DI benefits supply appears a robust determinant of both growth in DI rolls and declines in high school dropout labor force participation.

VII. Conclusion

Following the 1984 liberalization of the federal Disability Insurance program, the fraction of nonelderly adults receiving DI benefits rose by 60 percent. Contemporaneously, the DI recipient population became younger, less predominantly male, more likely to suffer from low mortality disorders, and far more encompassing of the low-skilled population. We trace these shifts to reduced screening stringency, rising cash income replacement rates, and the increasing in-kind value of Medicare benefits provided. We explore the impact of the supply of DI benefits on both the level of labor force participation among the less skilled and their employment responses to adverse demand shocks. As our model underscores, because disability application is more attractive to the unemployed—who intrinsically face lower opportunity costs of exiting the labor force to seek benefits—increases in the supply of DI benefits are likely to induce differential exit of low-skilled workers facing employment downturns, thereby reducing measured unemployment.

Using state level data from the Current Population Survey paired to SSA administrative data on disability applications, awards, and receipt by state and year, we implement two approaches to identify the labor market impacts of shifts in the supply of DI benefits. Initially, we exploit the interaction of state disability replacement rates with national changes in program stringency to identify plausibly exogenous state-level shifts in the

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36. Surprisingly, growth in the median wage for prime-age males is positively (i.e., wrong) signed in both the DI recipient and labor force participation regressions.

supply of DI benefits. This analysis finds that state level reductions in benefits supply induced large (relative) increases in the labor force participation of male and female high school dropouts during the DI retrenchment, followed by large declines during the DI subsequent expansion. No corresponding relationship is observed for changes in SSI rolls, or for changes in the labor force participation of more-educated workers.

We next investigate whether the declining stringency and growing generosity of the DI system have hastened the labor force exit of low-skilled workers facing adverse demand shocks. To distinguish supply from demand in this setting, we use cross-state variation in industrial composition to identify potentially exogenous shocks to employment of low-skilled workers both pre- and post-DI reform. Following the liberalization of DI in 1984, we find that DI application rates became two to three times as responsive to adverse labor demand shocks of equal magnitude. Simultaneously, male and female high school dropouts became almost twice as likely to exit the labor force in the event of an adverse shock—and correspondingly less likely to enter unemployment. While our reduced-form approach does not quantify the relative importance of reduced DI screening stringency and rising replacement rates in inducing these behavioral shifts, simple calculations suggest that the net impact of outward shifts in the supply of DI benefits is a plausible explanation for both phenomena.

Accounting for the role of the DI program in inducing labor force exit among the low-skilled unemployed, our illustrative calculations suggest that the measured U. S. unemployment rate among adults ages 25–64 in 1998 would have been about half a percentage point (13 percent) higher if it had not been for the liberalization of DI in 1984. Our analysis confirms the findings of Juhn [1992] and Juhn, Murphy, and Topel [1991] that declining real wages contributed to the labor force exit of the low skilled in this time period. But declining wages do not appear a sufficient explanation for our findings. Controlling for cross-state variation in the wage declines faced by low-skilled workers, we find that nonparticipation and disability receipt grew substantially more in states with larger shifts in the supply of DI benefits. In the absence of a significant change in DI program parameters, the aggregate size and accompanying labor market impact of the DI program is likely to grow substantially in the current decade.

<center>DATA APPENDIX</center>

A. 1. *Employment, Unemployment, and Labor Force Nonparticipation*

Annual, state-level data on employment, unemployment, and labor force nonparticipation 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 twelve 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*

We calculate two measures of the change in the opportunity wage faced by low-skilled workers in each state between 1984 and 1998: the change in the real fixed-weighted median log wage of high school dropout males ages 25–64 and the change in the real fixed-weighted log wage of prime age males 25–54 at the twenty-fifth percentile. Our sample includes all males in these age and education categories regardless of employment status excluding the self-employed (for whom wages are not given in the CPS), those working without pay, and employed workers for whom a wage could not be calculated (due to usual hours reported as "variable"). For workers employed in the survey reference week, wages were calculated by dividing usual weekly earnings by usual weekly hours. Observations were dropped for which calculated wages fell outside the range of $1 and $150 per hour (in year 2000 dollars). In cases where state-year-age-education cells were empty, we assigned the median value of corresponding nonempty cells. We use fixed age weights corresponding to the shares of males (male high school dropouts) in state population in 1984 ages 25–34, 35–44, 45–54, and 55–64. Wages are inflated to 2000 values using the chain weighted Personal Consumption Expenditure deflator.

A. 3. *State Level 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* [various years]. DI data includes only disabled workers receiving benefits whereas SSI data includes only disabled adult beneficiaries (thus excluding child and aged beneficiaries).

A. 4. *State Level DI and SSI Application and Award Data*

Administrative data on unique 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, the year with which disability applications and awards are associated is the year of application rather than the year of award.

A. 5. *Demographic Data on DI and SSI Recipients from the Survey of Income and Program Participation*

Since SSA does not report the educational distribution of DI recipients, we use the Survey of Income and Program Participation to estimate these numbers for 1984 and 1999. SIPP data are unfortunately not available for earlier years. Disability receipt rates by education category, age, and gender were estimated using data from the Survey of Income and Program Participation 1984 wave 1 and 1996 wave 12. A survey respondent is coded as a DI recipient if she "did receive income from Social Security for himself/herself in this month," and her reason for receipt of Social Security was disability. An individual is classified as an SSI recipient if she "did receive any income from Supplemental Security Income for him/herself during the reference period."

A. 6. *State Level DI Replacement Rates 1978 and National DI Replacement Rates 1979 and 1999*

To calculate 1978 state-level potential DI replacement rates used in Tables V, VIII, and IX, we use the CPS Social Security Earnings Records Exact Match file, which contains Social Security earnings records in each year 1951–1977 for currently employed males in the CPS sample. This is complete data for DI

benefits calculations. We calculate the potential DI replacement rate for all employed males ages 25–61 in each state in 1978 and use the seventy-fifth percentile in each state as a measure of the potential replacement rate faced by low wage workers. The 1978 CPS-SSER exact match file is used by Bound and Krueger [1991] and was helpfully provided to us by Alan Krueger.

For the replacement rate calculations in Table I, we utilize annual earnings data from the 1964–1999 March CPS Annual Demographic Files and the 1978 CPS Social Security Earnings Records Exact Match file (CPS-SSER). The March files provide annual earnings data for the years 1963–1998 and the Exact Match file is used to estimate earnings percentiles from 1951–1962. In each year between 1951 and 1998, we calculate the tenth, twenty-fifth, fiftieth, seventy-fifth, and ninetieth percentile of earnings for males at each age 25 to 61 inclusive to construct an age-year earnings file with 1776 observations (48 years, 37 ages). Because only earnings subject to the Social Security payroll tax are used in calculating Social Security benefits, all earnings percentiles in excess of the taxable maximum in a particular year are truncated at the maximum. We assume that an Nth percentile worker who was M years old in year T was an Nth percentile $M - t$ year old worker in year $T - t$ and estimate Average Indexed Monthly Earnings as follows:

(14)

$$AIME_{NMT} = \left(\frac{1}{\min[35, M-24]}\right) \cdot \sum_{m=\max[25, M-34]}^{M} \left(\frac{Y_{Nmt}}{12}\right) \cdot \max\left[\frac{\bar{Y}_{T-2}}{\bar{Y}_t}, 1\right].$$

In this equation, Y_{Nmt} is equal to taxable earnings for an m year old Nth percentile worker in year t where t is equal to $T - (M - m)$. The indexing series factor, $\max[\bar{Y}_{T-2}/\bar{Y}_t, 1]$, is obtained from the Social Security Administration and is equal to mean national earnings in year $T - 2$ divided by mean national earnings in year t (with a minimum value of 1). We use the median AIME within each of four age brackets (30–39, 40–49, 50–54, 55–61), to calculate potential DI benefits in year $T + 1$ by applying the PIA formula in equation (2). The Cash Income Replacement Rate given in Table I equals the ratio of DI benefits in year $T + 1$ to after-payroll-tax earnings in year T. Calculation for 1999 use the 1965–1999 March CPS files, while the 1979 calculations utilize both the 1964–1979 March CPS files and the 1978 CPS-SSER file (for earnings from 1951–1962). Because the Social Security Ad-

ministration only considered earnings after 1950 in calculating DI benefits in 1979, the maximum number of years of prior earnings in the 1979 AIME calculation is 28. Our calculation also assumes that individuals have a maximum of 35 years of earnings and that lowest earnings years occur before the age of 25. We do not impute wages for individuals who are out of the labor force. To the extent that male labor force nonparticipation has increased differentially at the low end of the earnings distribution, our algorithm is likely to understate the actual increase in replacement rates for low-skilled workers.

The simulated 1979 replacement rates found in Table I closely accord with the administrative figures from the CPS-SSER data. For example, the 1979 simulated replacement rates for males at the twenty-fifth percentile of the earnings distribution for age groups 30–39, 40–49, 50–54, and 55–61 are 41, 41, 41, and 45 percent. The corresponding replacement rates from the 1978 CPS-SSER data are 43, 41, 43, and 45 percent. We use the simulated replacement rates in 1979 for comparability with the simulated values for 1999 (for which no CPS-SSER data are available).

Replacement rates that include the value of Medicare benefits in the final two columns of Table I add average Medicare expenditures for DI beneficiaries in the relevant year to the DI benefit amount and scale earnings to account for average fringe benefit rates for individuals at each of the five earnings percentiles (tenth, twenty-fifth, fiftieth, seventy-fifth, and ninetieth) in 1982 and 1996, the two years closest to 1979 and 1999 that were available from Bureau of Labor Statistics data. Average Medicare expenditures for DI beneficiaries in 1979 were $2890 (in 1998 dollars) and in $4749 in 1998 (the most recent year for which data were available). Average benefit-wage ratios in 1982 were 0.0923, 0.1911, 0.2512, 0.2780, 0.2801 for the tenth, twenty-fifth, fiftieth, seventy-fifth, and ninetieth earnings percentiles, while the corresponding ratios in 1996 were 0.0778, 0.1743, 0.2554, 0.2965, 0.2988. Benefit-wage ratio data were kindly provided by Brooks Pierce of the Bureau of Labor Statistics and are based upon calculations in Pierce [2001].

A. 7. *Additional Variables Used in Section VI: Alternative Explanations*

a. The rate of nonaccidental, nonhomicidal mortality among state residents ages 25–64 is calculated using mortality data from U. S. Department of Health and Human Services [1984, 1998].

b. Nonelderly foreign-born populations by state for 1984 and 1998 are estimated using data from the 1980 and 1990 Census of Populations and the 1998 CPS. Prior to 1994, the CPS did not collect information on nativity, and hence we linearly interpolate 1984 state levels using the 1980 and 1990 Integrated Census Public Use Micro-samples (IPUMS) 1 percent files.

c. Incarceration rate data are calculated using Bureau of Justice Statistics data on state prison populations excluding local jails [U. S. Department of Justice 2002] denominated by state population ages 18–64 from the CPS.

d. UI generosity corresponds to the 1984 and 1999 potential state UI replacement rate for a full-time minimum wage worker. These data are obtained from the U. S. House of Representatives *Green Book* [Various years].

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