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CONSEQUENCES OF EMPLOYMENT PROTECTION? THE CASE OF THE AMERICANS WITH DISABILITIES ACT

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ABSTRACT

The Americans With Disabilities Act (ADA) requires employers to accommodate disabled workers and outlaws discrimination against the disabled in hiring, firing, and pay. Although the ADA was meant to increase employment of the disabled, it also increases costs for employers. The net theoretical impact turns on which provisions of the ADA are most important and how responsive firm entry and exit is to profits. Empirical results using the CPS suggest that the ADA had a negative effect on the employment of disabled men of all working ages and disabled women under age 40. The effects appear to be larger in medium size firms, possibly because small firms were exempt from the ADA. The effects are also larger in states where there have been more ADA-related discrimination charges. Estimates of effects on hiring and firing suggest the ADA reduced hiring of the disabled but did not affect separations. This weighs against a pure firing-costs interpretation of the ADA. Finally, there is little evidence of an impact on the nondisabled, suggesting that the adverse employment consequences of the ADA have been limited to the protected group.

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I. Introduction

Government efforts to eliminate employment and wage discrimination date back to the Equal Pay Act of 1963 and Title VII of the Civil Rights Act of 1964. These laws prohibited discrimination on the basis of race and sex. The consensus among labor economists is that civil rights legislation and related federal regulations led to a substantial improvement in economic conditions for Black Americans (see, for example, Freeman, 1981; Heckman and Payner, 1989; or Leonard, 1990), while the evidence on the impact of antidiscrimination policy on women is less clear cut (Leonard, 1989).

The most recent pieces of federal antidiscrimination legislation are the Americans with Disabilities Act (ADA) and the Civil Rights Act of 1991 (CRA-91). The ADA requires employers to offer reasonable accommodation to disabled employees, and bans discrimination against the disabled in wage determination, hiring, and firing. The ADA seems to be more far-reaching than CRA-91, which essentially modified existing antidiscrimination statutes.[1] ADA proponents hope the act will improve the labor market fortunes of disabled workers. Critics of the ADA have argued that adapting the workplace to the disabled can be expensive and that the costs of accommodation and ADA-related litigation may have significant negative employment effects (e.g., Rosen, 1991; Oi, 1991; Weaver, 1991; Epstein, 1992; Olson, 1997).

The first objective of this paper is to evaluate whether the ADA has in fact improved economic conditions for the disabled. A study of the ADA is also of broader interest, however. Although high-profile reasonable accommodation cases have attracted the most media attention, the majority of ADA charges are for wrongful termination. It is therefore possible that the ADA acts as a form of employment protection, similar to European firing costs. Since the ADA primarily affects a specific group, the consequences of employment protection may be easy to detect in this case. Moreover, contrary to early experience with the Civil Rights Act of 1964, in recent years, most discrimination charges under *all* statutes have been related to wrongful termination (EEOC, various years). If firing costs have disemployment effects, then this change in civil rights law enforcement may ultimately harm protected groups other than the disabled, a possibility raised by Donohue and Siegelman (1991, 1993).

The theory section of the paper uses a standard competitive model to highlight the distinction between hiring and firing costs due to the threat of lawsuits and the costs of accommodating disabled workers. Although the reasonable accommodation provision creates an incentive to employ fewer disabled workers, the introduction of hiring and firing costs complicates the analysis. If the threat of ADA-related litigation encourages employers to increase the hiring of the

[1] See Abram (1993). Oyer and Schaeffer (1998) explore the labor-market impact of CRA-91.

disabled, and if the number of employers is not very responsive to profits or costs, the ADA may increase the employment of disabled workers as ADA proponents had hoped. But when most charges are for wrongful termination and costs of reasonable accommodation are high, the ADA is likely to reduce disabled employment.

The empirical analysis looks at the employment and wages of disabled and nondisabled workers using data from the March Current Population Surveys (CPS) for 1988-1997. These data are useful for our purposes because the CPS income supplement identifies disabled workers, and because the March CPS has information on firm size, a variable that figures in our theoretical discussion, and in the ADA's compliance and sanction provisions. To investigate the impact of the ADA on turnover, we construct measures of separations and accessions by matching CPS rotation groups. Finally, we use EEOC data on discrimination charges by state to connect changes in labor market variables with the incidence of ADA-related charge activity.

The empirical results suggest that the ADA had a negative impact on the employment of disabled men, with no effect on wages. These results are unchanged by controlling for pre-ADA trends in disabled employment and for the increase in the fraction of people receiving Disability Insurance (DI) and Supplemental Security Income (SSI). In contrast to the results for men, the results for women are mixed. The ADA appears to have had a negative effect on women aged 21-39, but no effect on women aged 40-58. We also find no evidence that the ADA had a negative impact on nondisabled workers in any age group, suggesting that the adverse consequences of the legislation in this case have been limited to the protected group.

We check the basic findings in a number of ways. First, there is some evidence that employment of the disabled declined more in medium-size firms, possibly because small firms are exempt from the ADA and large firms can more easily absorb the costs imposed by the ADA. Second, effects are generally larger in states where there have been more ADA-related discrimination charges. Finally, there is some evidence of a negative impact on hiring of the disabled, but little evidence for a reduction in disabled separation rates. This suggests that the negative effects of the ADA may be due more to the costs of reasonable accommodation than to the threat of wrongful termination lawsuits.

The paper is organized as follows. Section II gives some background and discusses related literature. Section III provides a theoretical analysis of provisions that protect disabled workers. Section IV describes the data and our empirical strategy. Section V contains the main empirical findings and Section VI concludes.

II. Background

A. ADA Provisions and Coverage

The ADA was signed into law in July 1990 and came into effect in July 1992. Previously, there was no federal law dealing with the employment and wages of disabled workers in the private sector, although the Rehabilitation Act of 1973 covered disabled workers employed by the federal government or working for federal contractors. A number of states also had laws protecting disabled workers, but the coverage and effectiveness of these laws varied. Title I of the ADA initially covered all employers with at least twenty five employees. In 1994, coverage was extended to employers with fifteen or more employees. Title I requires employers to provide "reasonable accommodation" for their disabled workers. Examples include enabling wheel chair access, purchasing special equipment for disabled employees, and job restructuring to permit disabled employees to work part time or from home. Title I also bans discrimination against the disabled in wages, hiring, firing, and promotion. For example, a disabled employee should be paid the same amount as a nondisabled worker in the same job, and firms are not allowed to consider disability in hiring and firing decisions.[2]

Enforcement of ADA provisions is left to the Equal Employment Opportunity Commission (EEOC) and the courts. Disabled employees or job applicants who believe they have been discriminated against can file a charge with the EEOC, which will investigate and in some cases try to resolve the charge or sue. If the charge is not resolved and the EEOC does not sue on behalf of the charging party, it issues a letter of permission to sue and the charging party is free to litigate at his or her own expense. The law provides for remedies that include hiring, reinstatement, promotion, back pay, front pay, and reasonable accommodation, and for payment of attorney's fees, expert witness fees, and court costs. As a consequence of CRA-91, compensatory and punitive damages are also available if intentional discrimination is found. These range from $50,000 for firms with 100 or fewer employees to $300,000 for firms with 500 or more employees (EEOC, 1995, p. X-8).

The underlying logic of the ADA seems to be that employers incorrectly perceive the disabled to be less productive, or are unwilling to make modest adjustments to accommodate them (see, e.g. Kemp, 1991). The fact that the labor market fortunes of the disabled are much worse than the nondisabled is not in dispute. The disabled earn 40 percent or less of what nondisabled workers earn. Their labor force participation rates are much lower and they are much less likely to be employed (see, e.g., Burkhauser and Daly, 1996, or our statistics,

[2]Title II covers discrimination in public programs, and Titles III and IV refer to public accommodations (businesses) and telecommunication. Title V contains technical information related to enforcement (see EEOC, 1995).

below). ADA proponents believe the law will induce companies to make the investments and modifications necessary to employ disabled workers, and reduce unjustified discrimination. In recent years, interest in the labor market performance of the disabled has also been fueled by efforts to reduce the number of Disability Insurance recipients (see, e.g., Leonard, 1991)

From July 1992 to September 1997, the EEOC received 90,803 ADA charges (EEOC, various years). Of these, 29 percent mention "failure to provide accommodation", while 9.4 percent are for discrimination at the hiring stage. The majority of charges, 62.9 percent, are for wrongful termination (i.e., discharge, failure to rehire, suspension, or layoff). This motivates our interpretation of the ADA as providing a form of employment protection.

B. How Costly is the ADA?

We have not found representative data on the costs of accommodation, though the President's Committee on Employment of People with Disabilities has surveyed some employers who contacted them for help accommodating their disabled workers (Job Accommodation Network, 1997). This survey shows an average cost of $930 per accommodation since October 1992. Other costs of accommodation include time employers spend dealing with ADA regulations and reduced efficiency due to a forced restructuring.

An important component of ADA costs results from the threat of litigation. Since July 1992, over 11,000 of the charges brought under the ADA were resolved by the EEOC, and employers paid over $174 million in settlements (EEOC, various years). This figure does not reflect administrative costs, lawyers fees, and private settlements in or out of court. Although we do not have data on ADA suits alone, Condon and Zolna (1997) report that employees file over 40,000 cases each year with state and federal courts, the majority related to discrimination, and win almost 60 percent of the time. They estimated an average award of over $167,000 and defence costs of over $40,000 (less than the $80,000 estimated by Dertouzos, 1988, for wrongful termination suits in California).

The ADA may also have been a factor in the development of a new insurance market, policies for Employment Practices Liability Insurance (EPLI), which covers the costs of employee lawsuits. The EPLI market started in late 1990 and has since grown rapidly, with minimum premia ranging from $4,500 to $20,000 a year (Clarke, 1996). This suggests that the costs of the ADA are indeed a real concern for employers.

C. Related Literature

DeLeire (1997) is the first systematic empirical study of the impact of the ADA that we are aware of. He used the Survey of Income and Program Participation and the Panel Survey of Income Dynamics to compare labor market outcomes for disabled and non-disabled workers before and after the ADA. Our approach uses different data and a number of empirical strategies not explored by DeLeire, but we also begin with similar comparisons of changes in outcomes by disability status. Since the ADA creates firing costs, our paper is also related to empirical work on the impact of firing costs in Europe (e.g., Lazear, 1990; Addison and Grosso, 1996; and Nickell, 1997). Related theoretical papers include analyses of labor market regulations by Mortensen (1978), Summers (1989), Lazear (1990), Bertola and Bentolila (1990), and Hopenhayn and Rogerson (1993), and the analysis of optimal disability benefits by Diamond and Sheshinski (1995). Finally, our work is also related to the empirical literature on the impact of mandated benefits (e.g., Gruber, 1994; Ruhm, 1996; Waldfogel, 1996) and, as noted above, to analyses of the impact of earlier civil rights legislation.

III. Consequences of Protecting Disabled Workers: Theory

The theoretical consequences of the ADA are explored using a standard competitive model with two types of workers. Nondisabled workers supply labor according to the function $n_a(w_a)$ and the disabled supply labor according to $n_d(w_d)$, where w_a is the wage received by nondisabled workers and w_d is the wage rate for disabled workers. We assume that n_i is increasing in the wage rate for $i = a, d$. All workers are infinitely lived and risk-neutral, and have a discount factor $\beta < 1$. There are M firms in the labor market which never exit, and a large number of potential firms that can enter at cost Γ. This is a convenient formulation enabling us to discuss both a market characterized by free entry (when $M \to 0$) and one where the number of firms is fixed ($M > 0$ and $\Gamma \to \infty$).

All firms are risk-neutral, discount the future at rate β, and have access to the production function $F(L_t, D_t)$, where L_t is the number of nondisabled workers and D_t the number of disabled workers employed at time t. We also assume that with probability s every period, the productivity of a worker at his or her current firm falls to zero, though productivity elsewhere is unaffected (this may be due, for example, to match-specific learning as in Jovanovic, 1979).

We assume $F(L_t, D_t) = \frac{1}{\alpha}\left(L_t^\rho + eD_t^\rho\right)^{\frac{\alpha}{\rho}}$, where $\alpha < 1$, and $0 < \rho \leq 1$. The quantities L_t and D_t include only employees who have not received adverse match specific shocks. The parameter e captures the relative productivity of disabled workers. For example, when $\rho = 1$ and $e < 1$, disabled and nondisabled workers are perfect substitutes, but the disabled have a lower marginal product than nondisabled workers. This formulation also nests the case in which firms discriminate against the disabled for taste reasons as in Becker (1971).

The issue raised most often in ADA charges is wrongful termination, followed by failure to provide reasonable accommodation and discrimination in hiring. Suppose disabled job applicants who are not hired sue with probability p_d at expected cost ν_d, including damages and lawyer fees. Rejected nondisabled applicants can also sue, falsely claiming to be disabled, and this happens with probability p_a and has cost ν_a. The expected cost of not hiring a disabled worker is therefore $h_d = p_d \nu_d$, while the corresponding cost for a nondisabled worker is $h_a = p_a \nu_a$. We refer to h_a and h_d as hiring costs, though more precisely, these are costs that the firm incurs when it decides not to hire an applicant. A disabled worker who is fired sues with probability q_d for damages ϕ_d. For a nondisabled worker, the corresponding probabilities and damages are q_a and ϕ_a, so the expected costs of firing a disabled and nondisabled worker are $f_d = q_d \phi_d$ and $f_a = q_a \phi_a$. We begin with the simple case where all costs are pure waste, so h and f act like a tax from the point of view of the worker and the firm, though suits may benefit some other party such as lawyers and insurance companies. This seems like a reasonable starting place since a significant fraction of the litigation costs imposed on employers probably do not get transferred to disabled workers. Obviously, a fraction of these costs do accrue to workers, giving them a reason to sue, but we defer a discussion of this case.

We assume that $(1-\beta) f_a < w_a$ and $(1-\beta) f_d < w_d$, which implies that firms always wants to lay off the fraction s of its employees who receive adverse match-specific shocks. We also assume there is an excess number of applicants for every job, D_F of whom are disabled, and L_F are nondisabled. We treat D_F and L_F as given. Finally, we assume that firms can "accommodate" disabled workers, for example by purchasing special equipment at cost C per worker, and that this expenditure increases the marginal productivity of disabled workers by a fixed amount B per worker. The ADA requires employers to make such accommodations. If $C < B$, employers would make these adjustments voluntarily, even in the absence of the ADA. The fact that government regulation is required suggests that typically $C > B$.

The maximization problem of the firm at time $t = 0$ can be written as:

$$\begin{aligned} \max_{\{D_t, L_t\}} \Pi \equiv & \sum_{t=0}^{\infty} \beta^t \left[F(L_t, D_t) - w_{a,t} L_t - w_{d,t} D_t - c D_t - f_a s L_{t-1} - f_d s D_{t-1} \right. \\ & \left. - h_a \left(L_F - [L_t - (1-s) L_{t-1}] \right) - h_d \left(D_F - [D_t - (1-s) D_{t-1}] \right) \right] \end{aligned} \quad (1)$$

where $L_{-1} = D_{-1} = 0$, $w_{a,t}$ and $w_{d,t}$ denote nondisabled and disabled wages at time t, and $c = C - B$ is the net cost of accommodation after the ADA. Pre-ADA firing and hiring costs and costs of accommodation are assumed to be zero. The first line of (1) is output minus the wage bill, accommodation costs and firing costs. Firms discharge a fraction s of their employees who receive an adverse match-specific shock, incurring a firing cost of f_d for each disabled layoff, and f_a for every nondisabled termination. The second line gives the "hiring costs" the firm incurs as a function of the number workers not hired out of the applicant pools, L_F and D_F. When $L_t = L_{t-1}$ and $D_t = D_{t-1}$ so that employment is not changing, the firm hires sL_{t-1} nondisabled and sD_{t-1} workers to replace those who are laid off. As noted above, h_a and h_d act as *hiring subsidies*, because the firm reduces its costs by hiring more workers.

Since adjustment costs are linear and there is no aggregate uncertainty, firms immediately adjust to steady state employment levels, and $w_{a,t} = w_a$, $w_{d,t} = w_d$, $L_t = L$, and $D_t = D$ in every period. These equilibrium employment and wage levels satisfy:

$$\begin{aligned} L^{\rho-1}\left[L^{\rho}+eD^{\rho}\right]^{\frac{\alpha-\rho}{\rho}} &= w_a+\beta s f_a-\left[1+\beta(1-s)\right]h_a \\ eD^{\rho-1}\left[L^{\rho}+eD^{\rho}\right]^{\frac{\alpha-\rho}{\rho}} &= w_d+\beta s f_d-\left[1+\beta(1-s)\right]h_d+c. \end{aligned} \tag{2}$$

Both equations equate the relevant marginal product to the flow marginal cost, inclusive of firing costs, hiring subsidies, and the net costs of accommodation..

To determine equilibrium, we also impose market clearing for nondisabled workers,

$$n_a^{-1}\left(mL\right) = w_a, \tag{3}$$

where n_a^{-1} is the inverse supply function and m is the equilibrium number of firms. The number of firms is determined by

$$\Pi \leq \Gamma \text{ and } m \geq M, \tag{4}$$

which holds with complementary slackness. This means that either the maximized value of profits is equal to entry costs, or that there is no entry and the number of firms, m, is equal to the minimum, M.

Finally, the wages received by disabled workers are given by

$$w_d = \max\left[n_d^{-1}\left(mD\right), \eta w_a\right], \tag{5}$$

where η is a parameter. In the absence of the ADA, $\eta = 0$ and there are no restrictions on disabled wages. The equal pay provision can be interpreted as setting $\eta = 1$.

Inspection of the equilibrium conditions yields the following conclusions:

1) From (2), it is clearly possible for the ADA to reduce the costs of employing the disabled, since h_d is effectively a hiring subsidy. The scenario envisaged by ADA proponents can probably be best described as $h_d > 0$, $f_a = h_a = 0$, and minimal firing costs for the disabled, in which case the ADA can increase employment and wages of the disabled. It is also worth noting that

as long as $\rho > 0$, the ADA will reduce nondisabled employment and wages in this case, though this effect is likely to be small.

2) As we noted in Section II, the ADA appears to have increased f_d considerably more than h_d. Similarly, costs of employing the disabled are increased by the accommodation costs, c. Therefore, in practice, the ADA may be more likely to reduce disabled employment and wages.

3) The equal-pay provision of the ADA (i.e. $\eta > 0$) may have increased disabled wages, creating involuntary unemployment off the disabled supply curve. The equal-pay provision also interacts with firing costs and the costs of accommodation by preventing wages from falling to offset these costs, exacerbating the decline in disabled employment.

4) Although the partial-equilibrium effect of hiring costs is to increase employment of the disabled, the implicit hiring subsidy h_d is effectively financed by reducing profits. If $m > M$ and $\Pi = \Gamma$ to start with, then an increase in h_d will cause some firms to shut down, causing both disabled and nondisabled employment and wages to fall. More generally, the contrast between the free-entry and fixed-number cases suggests that the ADA will reduce employment most in firms or industries where profits are close to entry costs. In the empirical work, we use firm size as a proxy for excess profitability, since large firms are more profitable (see, e.g., Schmalensee, 1989; Scherer and Ross, 1990, Chapter 11).

5) Finally, the ADA might also affect employment by changing firing and hiring costs for the nondisabled, f_a and h_a.

The discussion so far presumes that f and h act like taxes on the firm-worker relationship rather than a transfer from the firm to the worker. We know from the work of Mortensen (1978) and Lazear (1990) that when f and h are pure transfers and side payments are allowed, firing costs should have no effect on employment. To see this in our context, suppose that $f_a = h_a = h_d = 0$ and $f_d > 0$. Under the assumption that f_d is a pure transfer and both parties are risk-neutral, labor supply of disabled workers changes to $n_d(w_d + \beta s f_d)$. The reason for the change is that workers anticipate they may be fired, and therefore include the discounted flow value of firing costs, $\beta s f_d$, in their employment income. It is straightforward to see that as long as w_d can fall and keep $w_d + \beta s f_d$ constant, changes in f_d have no effect on disabled employment. However, the equal-pay provision of the ADA limits this possibility. Moreover, because f and h may include payments to third parties and because firms and workers are risk-averse, characterizing these costs as a tax on the employment contract seems like a better stylized description of reality.

Finally, we note that the analysis so far ignores the impact of firing costs on separations because separations are exogenous. In a previous version of the paper (details available upon request), we allow for time-varying productivity and endogenous separations. This analysis shows that ADA-related firing costs are likely to reduce both hiring and separations.

The theoretical discussion therefore shows that the net effect of the ADA depends on which provisions are most important. The accommodation and firing costs imposed by the ADA are likely to reduce employment, while hiring costs have the opposite effect. If the equal-pay provision is not binding, equilibrium is on the labor supply curves of disabled and nondisabled workers, and employment declines are accompanied by wage declines. More generally, however, the equal-pay provision can create "involuntary unemployment" off the supply curve. We therefore estimate reduced-form equations of the form:

$$y_{it} = x'_{it}\beta + \delta_t + \gamma_t a_{it} + \varepsilon_{it}, \tag{6}$$

where i denotes individuals and t time; y_{it} is weeks worked or average weekly wages; and x_{it} is a set of controls including disability status. The term δ_t is a year effect to control for changes in aggregate conditions. Finally, a_{it} is the interaction of a disability status indicator and a dummy for post-ADA years. The coefficient γ_t measures the impact of the ADA on the disabled using the nondisabled as a control group. Since the ADA potentially affects the employment of nondisabled as well as disabled workers, we also explore empirical specifications that use variation by firm size and state to separately identify effects on the disabled and nondisabled.

IV. Data and Descriptive Statistics

The sample is drawn from the March CPS's from 1988 through 1997, and limited to men and women aged 21-58, since this age group has strong labor-force attachment. Disabled workers are identified in the March CPS Income Supplement by the question: "Do you have a health problem or disability which prevents you from working or which limits the kind and the amount of work you can do?". This question has been used by other researchers working on disability issues (e.g., Krueger and Kruse, 1995), and is similar to disability questions in the PSID and SIPP (see, e.g., Burkháuser and Daly, 1996; DeLeire, 1997).[3]

The impact of the ADA on employment levels is evaluated by looking at data on weeks worked during the previous calendar year from the March income supplement. The wage measure is average weekly earnings, computed using annual earnings data from the supplement.

[3]Using data from the Retirement History Survey and the National Longitudinal Survey of Older men, Bound (1989) compares objective measures of health status with self-reported measures like the one used here. His results suggest there is no clear basis for preferring one type of measure over the other.

Although the CPS changed from paper questionnaires to computer-assisted interviewing in 1994, and the main labor force status questions were also revised at that time, the content of the income supplement was not changed. Appendix A discusses an analysis of matched CPS data for March 1993 and 1994, and provides more information on the CPS redesign and its possible consequences for our analysis.[4]

The variables in the income supplement refer to the previous calendar year, so the sample has data for weeks worked and wages in 1987-1996. The disability status question in the supplement appears to refer to respondents' status at the time of the survey (March of the survey year), but actually serves as a lead-in question prefacing additional supplement questions about disability income the previous year. Except for Table I and Figure I, which present descriptive statistics dated by survey year, the tables and figures label estimates according to the year of observation, which is the survey year minus one.

CPS supplement respondents provide information on the size of the employer they worked for longest in the past year. Responses to this question are grouped into three brackets; 1-24 employees (small), 25-99 employees (medium), and 100 or more employees (large). The analysis of hiring and firing outcomes is based on changes in employment status in a match of CPS rotation groups from March to April. Finally, we use EEOC data on ADA-related charges brought in each state since 1992 to see if charge activity is related to employment.[5]

Descriptive statistics organized by survey year, age group, and sex, are reported in Table I. The table shows no major change in average weeks worked by men or in the proportion of men currently employed, although there is a modest increase in weeks worked by women. Table I and Figure 1 show an increase in work-related disability rates for men and women. Disability rates for men aged 40-58 start increasing in the 1991 survey, and remain high thereafter, though with a slight decline in 1996 and 1997. Disability rates for women aged 40-58 increase sharply starting in 1994. For men aged 21-39, there is a small increase in self-reported disability rates between 1990 and 1994, which is later reversed, and in 1996 the disability rate for this group is lower than it was in 1990. These patterns suggest the ADA may have had an effect on the probability that people describe themselves as disabled, especially for women.[6] This raises the

[4]The estimates we report are weighted by CPS sample weights. The weights were updated in 1994 to reflect the 1990 Census. We use newly released updated weights for the 1990-1993 surveys as well. The 1988 CPS data come from the so-called March "rewrite" file. This file includes firm size and other variables not on the original 1988 release, and reflects a revised imputation procedure (Bureau of the Census, 1991). The extract excludes the hispanic oversample for each year. A few dozen households with duplicate household identifiers in the 1994 survey were also excluded because they could not be included in the matched samples.

[5]The analysis variables are state/year aggregates from our tabulations of EEOC microdata. These are similar to less recent statistics published in EEOC reports.

[6]The fact that the disabled may be an elastic population has been noted by Oi (1991) and Kubik (1997). On the other hand, Dwyer and Mitchell (1998) argue that disability status does not appear to be endogenous in models of retirement behavior.

possibility of composition effects, a point we return to in the discussion of results.

V. Results

A. Employment and Wage Effects

Figure 2 plots average weeks worked by sex and disability status. While men aged 21-58 work on average over 45 weeks a year, disabled men work less than 20 weeks on average. Especially noteworthy is the 2 week decline in weeks worked by disabled men between 1992 and 1993, the first full year in which the ADA was in effect. There is no sharp decline in weeks worked by disabled women aged 21-58.

Figures 3a and 3b plot average weeks worked separately by age group. Weeks worked by disabled men aged 21-39 dropped sharply between 1992 and 1993, while those by disabled women aged 21-39 started falling in 1992. Weeks worked by men aged 40-58 also show a marked decline between 1992 and 1993. In contrast, there was an increase in weeks worked by disabled women aged 40-58 between 1992 and 1993, which is why we do not see a decline for disabled women when the two age groups for women are pooled.

Table II reports ordinary least squares (OLS) estimates of equation (6). These estimates are from regressions of weeks worked and log weekly wages on age and race dummies, year and disability effects, and year×disability interactions for 1991-1996. The interaction terms are the coefficients of interest. Because the ADA came into effect in July 1992, we think of 1993-96 as post-treatment years. 1992 is possibly a transition year, while the effects for 1991 can be seen as "pre-treatment" specification tests. The table also reports estimates from specifications including a linear time trend interacted with disability status, and a single-effect model that includes a trend interaction with disability status, 1991 and 1992 interactions with disability status, and a single post-ADA dummy for 1993-1996 interacted with disability status. The models with trends allow for the possibility that changes in outcomes by disability status can be explained by extrapolating different trends for the disabled and nondisabled.

The results in Table II suggest that the ADA had a substantial and statistically significant negative effect on the employment of disabled people under 40. For example, column (1) of Panel A shows that weeks worked by disabled men aged 21-39 were stable until 1992, but fell by 1.8 weeks in 1993, and declined by an additional week in 1995. Column (2) shows that controlling for disability-specific trends does not change these results. Results for women aged 21-39 are similar (columns 4-6), but the decline starts in 1992. The sharp employment declines in 1992 and 1993 suggest that the costs of the ADA were not anticipated by employers before the law became effective. This is not too surprising since the first years of ADA enforcement were characterized by some confusion regarding exactly what the act required (see, e.g., Veres and Sims, 1995).

Panel B reports estimates for the 40-58 age group. The disability-year interaction for men is 1.8 weeks in 1993, but this effect is halved when a linear trend is added. On the other hand, the single-effect model reported in column (3) generates a statistically significant decline of 2.1 weeks after controlling for disability-specific trends. Finally, column (4) shows a decline in the relative employment of disabled women aged 40-58, but this is in 1991, before the ADA came into effect. Moreover, columns (5) and (6) show that these effects disappear once we control for disability-specific trends, and, in fact, the coefficients of interest change sign. So there is little evidence that the ADA had an effect on the employment of disabled women in the 40-58 age group.

Columns (7)-(12) in Panel A report estimates for the log weekly earnings of men and women aged 21-39. There is an effect on men in 1994, but not in other years, and the 1994 effects disappears in models with a trend. There is no evidence of a wage effect for women aged 21-39. The wage estimates for older men, reported in Panel B, are also sensitive to the inclusion of trends. Moreover, like the employment effects for older women in Panel B, the decline in wages for older women starts too soon. We therefore interpret the wage results as suggesting that the ADA had little effect on the relative wages of disabled workers. The rest of the paper focuses on further analyses of employment effects only, and the analysis is limited to demographic groups where the evidence for effects is strongest—women aged 21-39, and men in both age groups.

Composition Bias

A possible explanation for the employment results in Table II is a composition effect. Figure 1 shows an increase in self-reported disability rates after 1991. If unemployed people were disproportionately more likely to identify themselves as disabled after the ADA, perhaps because disability became more socially acceptable, the results in Table II could overestimate the disemployment effects of the ADA.

We investigate the possibility of composition bias using a matched sample from the 1993 and 1994 CPS's. In principle, the matched sample provides two observations for half of the 1993 respondents. (In practice the match rate is lower; see Appendix A for details). The matched sample is used here to compare individuals who report a disability in both surveys to those who do not report a disability in either year. Since these surveys report data on weeks worked in 1992 and 1993, the matched data set provides a short panel that straddles the ADA's implementation date and is unaffected by composition bias arising from changes in reporting behavior.

Results using the matched sample are reported in Appendix Table A1. Column (1) repeats the basic cross-sectional specification for 1993 and 1994 data only. Column (2) reports results for the sample of individuals for whom the match was successful. Column (3) reports results for the sample of individuals reporting the same disability status in both years. The cross-sectional estimates in Table A1 are similar to those in Table II. Comparing columns (1) and (3), it is apparent that the restriction to the same disability status in both years has almost no effect on the results for men aged 40-58 and women aged 21-39. This is important evidence against the presence of composition effects since disability rates increased more for these two groups than for the younger men.

For men aged 21-39, the results using a sample where disability status is unchanged are smaller than in the 93-94 cross-section. The disability-constant sample shows a 0.3 decline in weeks worked, as compared to a 1.2 week decline in the full cross-section, or a 3.1 week decline in the matched cross-section. Adding controls for personal characteristics changes this to a 1 week decline, which is closer to the 1.5 week decline observed in the corresponding full cross-section. Finally, excluding observations where supplement and weeks worked data were allocated by the Census Bureau generates an estimate of -1.3 (reported in column (7)). Overall, these estimates suggest that the results for men aged 21-39 are not generated by composition effects either. In fact, the composition story is not plausible for this demographic group anyway, since disability rates for men aged 21-39 are actually lower at the end of the sample period than at the beginning, and the increase in disability rates could account for at most half of the decline in weeks worked.

Changes in the SSI and DI Programs

Another factor affecting the interpretation of Table II is an increase in the number of people receiving disability payments from the Disability Insurance (DI) and Supplemental Security Income (SSI) programs in the early 1990s (see, e.g., Stapleton, *et al*, 1994). Disabled workers who worked long enough are entitled to receive DI payments when not engaged in substantial gainful activity. Disabled people without a work history can receive SSI, which is a means-tested federal benefit that is supplemented by some states. Since SSI and DI payments create adverse labor supply effects, increased use of these programs may account for the decline in disabled employment (a possibility suggested by Weidenbaum, 1994). We address this issue in Appendix B using CPS data on social security program use. The results of this analysis suggest that the results in Table II are not explained by trends in SSI and DI receipt (see Table B1 for the estimates).

Magnitudes

The estimates in Table II can be compared to estimates of the effect of the ADA on the costs of employing disabled workers. Unfortunately, there are no good estimates of these costs, so our calculations are really just educated guesses. The average EEOC charge rate was about 12 per 1000 disabled employees a year. In cases resolved by the EEOC, about 14 percent of all ADA charges, employers made average payments of over $15,000 per case. In the remaining cases, either the charge was dropped, is pending, or there was a suit. We do not know what fraction of ADA charges end up in court. However, between 1995-97 there were over 56,000 employment discrimination cases brought in federal court (Administrative Offices of U.S. Courts, 1997). The total number of employment discrimination charges filed with the EEOC during this period was 245,000, which implies that 23 percent of all charges went to court. We apply this fraction to ADA charges, and use the cost estimate of $210,000 per case offered by Condon and Zolna (1997). To be on the conservative side, we also assume that if an ADA charge does not go to court or get settled by the EEOC, there are no other costs. The estimated average cost of an ADA charge is therefore equal to $0.23 \times 210,000 + 0.14 \times 15,000 = 50,400$. This is $50,400 \times 0.012 =$ $605 a year per disabled employee, or $12 for each week of exposure to the risk of a suit.

Assessing the cost of reasonable accommodation is even harder. The Job Accommodation Network (1997) reports a monetary cost of $930 per accommodation, which we take as the net cost.[7] Our estimates of separation rates in the next section suggest that the average duration of a job held by a disabled employee is 10 months, which implies that accommodation leads to a $23 increase in weekly employment costs. The average weekly earnings of the disabled were about $365 in 1991 and 1992. Since the total weekly cost increase due to the ADA is about 12+23=35 dollars, our rough estimate is that the ADA led to a 10 percent increase in the cost of employing disabled workers.

In the theoretical model in Section III, employers take the total cost of labor as given and are always on their labor demand curve. Since the results in Table II show little evidence of a change in disabled wages, the cost increase generated by the ADA falls on employers. The 10 to15 percent decline in weeks worked is therefore consistent with demand elasticities of about -1 to -1.5 for disabled workers. This is in the range of elasticity estimates reported by Hammermesh (1986) for workers in different demographic groups.

[7] This number excludes any productivity increases due to accommodation and any losses from job changes or sub-optimal reorganization of the work environment.

B. The Impact of the ADA on Hiring and Separations

We used matched CPS rotation groups from March to April to investigate the effect of the ADA on hiring and separation rates (see Appendix A for details relating to the March-to-April match). An individual is coded as having experienced a separation in year t if he or she is employed in March of that year and not in April. Similarly, an accession (hire) is recorded when someone who was not employed in March is employed in April. Separations are defined for those working in March while accessions are defined for those not working in March. Disability status always refers to March. These measures of accession and separations are the same as those used by Poterba and Summers (1986), and the resulting average accession and separation rates are close to those they report.[8]

The estimates of effects on separations and accessions are imprecise and also potentially affected by the CPS redesign (since the underlying data come from the main CPS survey and not the supplement). Moreover, Poterba and Summers (1986) show that labor-market transition data are plagued by considerable measurement error. We therefore limit the discussion in this section to a brief graphical analysis.

Figure 4a-d plot log accession and separation rates by disability status, sex and age group.[9] Figure 4a shows a post-ADA decline in separation rates for disabled workers aged 21-39, though there is also a decline for the nondisabled. On the other hand, Figure 4b shows a post-ADA decline in the hiring of disabled workers aged 21-39 that is not mirrored in the data for the nondisabled, especially for women. Figure 4c, which plots separation rates for the 40-58 age group, again shows no evidence of changes unique to the disabled. On the other hand, while not as clear as the change in disabled hiring for the young group, Figure 4d shows some evidence of a relative decline in hiring for disabled men in the older group. The apparent reductions in hiring for men and younger women are not surprising since employment for these groups fell. It is important to note, however, that the lack of a clear reduction in disabled separations weighs against a pure "firing costs" model of the ADA.

C. Results By Firm Size

Next, we look at employment patterns by firm size. This is of interest because firms with less than 15 employees are not covered by the ADA, and those with 16-25 employees were

[8]The average separation rate for 1988-1997 is 0.026 for men and 0.036 for women. The average accession rate is 0.18 for men and 0.08 for women. The disabled have higher separation rates and lower accession rates than the nondisabled.

[9]For instantanous accession and separation rates, η and ζ, the steady-state employment rate is approximately $e = \frac{\eta}{\eta+\zeta}$. We plot log accession and separation rates because $\frac{de}{dx} = \left(\frac{d\log\eta}{dx} - \frac{d\log\zeta}{dx}\right) \cdot (e(1-e))$.

initially exempt. The ADA might also have had a larger effect on employment in small firms since, as noted in Section III, small firms are probably less able to absorb ADA-related costs. Together, these considerations suggest we might expect the ADA to have had the largest effect on employment in firms that are sufficiently large to be covered by ADA provisions but small enough to be vulnerable to an increase in costs.

Figures 5a-c plot the log of the probability of working in a particular firm size category divided by the probability of not working. As noted earlier, the size category refers to the worker's longest job last year as recorded in the supplement. The size categories are 1-24 (small), 25-99 (medium), and 100 or more (large). The figures give a visual representation of the coefficients in a multinomial logit model where the dependent variable is employment-by-size-category, and non-workers are the reference group. The independent variables are year effects. The log-odds in each figure were computed separately for disabled and non-disabled workers.

The log-odds of working in a medium size firm appear to have fallen somewhat more steeply than log-odds of working in a small firm after 1992 for disabled workers in all three demographic groups. For women, there is also a relative decline in the probability of working in a large firm. Estimates of these differing trends by firm size are not very precise (e.g., $t=1.4$ for the medium vs. small contrast for men aged 21-39), but they are negative for all three demographic groups. In contrast with this pattern, the log-odds of employment by firm size are essentially parallel for nondisabled workers, suggesting the ADA had no effect on the nondisabled. Of course, even if there were effects on the nondisabled, it seems likely that they would be much smaller than effects on the disabled, and therefore harder to detect.

D. Cross-state Variation in ADA Charge Rates

The last strategy we use to estimate the impact of the ADA relates changes in employment to state-level variation in ADA charge rates. Like the firm-size analysis, this strategy allows us to separately identify the impact of the ADA on disabled and nondisabled employment. For the purposes of this analysis, the ADA charge rate is defined as the number of ADA-related discrimination charges.in a state, per thousand population disabled in the state in 1992 (the latter figure estimated using the 1992 CPS). Charge rates are calculated separately for 1992, 1993, 1994 and 1995, but 1992 population data are always used for the denominator.

Charge rates vary considerably by state. The average cumulative rate for 1993-95 was 13 per 1000 disabled persons aged 21-58, varying from a minimum of 6/1000 to a maximum of 40/1000. Variation in charge rates is generated by idiosyncratic differences in state labor force composition, local awareness of ADA provisions, cross-state differences in employers' compliance with the ADA, and whether a state previously had an FEP statute that covered disabled workers. Some states had weak laws, while others had laws that set criminal as well as civil penalties in cases where discrimination is proved (Advisory Commission on Intergovernmental Relations, 1989).

The estimating equation for this strategy is:

$$y_{idst} = x_i'\beta + \phi_{ds} + \delta_{dt} + \sum_{\tau=93}^{96} \alpha_{d,\tau}(f_{i,\tau} r_{s,\tau-1}) + \eta_{idst}, \tag{7}$$

where y_{idst} is weeks worked by individual i with disability status d living in state s in year t, and x_i is a vector of individual characteristics (age and race dummies). The regressors of interest are interactions between year dummies (e.g., $f_{i,93}$) and the charge rate per disabled person in the *previous* year in the individual's state of residence (e.g., $r_{s,92}$). Equation (7) also includes separate state effects for each disability group, ϕ_{ds}, and year×disability effects, δ_{dt}. The parameters of interest, $\alpha_{d,93}$, $\alpha_{d,94}$, $\alpha_{d,95}$ and $\alpha_{d,96}$, are the coefficients on the state/year charge rate interactions for each disability group. These effects are identified because we do not have a full set of state×year×disability interactions in the models. The year×disability effects now play the role of control variables. We also report estimates from a model that adds linear trends for each state×disability group to equation (7).

The results in Table III show that in 1993-94, weeks worked by the disabled men in states with a large number of ADA charges declined relative to other states. Some of the estimates for women aged 21-39 and men aged 40-58 are negative and significant in other years as well. The effects in 1993 are implausibly large, probably because the charge rate in 1992 was very low, but estimates for other years are of a plausible magnitude. For example, halving the mean annual charge rate of 4/1000 is predicted to reduce employment by 2.6 weeks based on the 1994 estimate for men aged 21-39.

In contrast with the estimates for the disabled, there is no consistent evidence of a negative impact of ADA charges on the non-disabled. The only significant negative estimates for the nondisabled appear in models that control for linear trends. For the most part, these estimates are considerably smaller than the corresponding estimates for the disabled.

As a final strategy, we experimented with IV estimation of (7) treating $d_{i,\tau} r_{s,\tau-1}$ as endogenous. The instrument was a dummy for whether a state previously had an FEP law restricting discrimination against the disabled, interacted with disability status and post-ADA year dummies. The presence of a preexisting FEP law is negatively correlated with ADA charge rates, probably because the ADA was less of an innovation in states with their own laws restricting discrimination against the disabled. For example, the states with preexisting FEP laws had

15 percent lower ADA charge rates. The resulting IV estimates are imprecise but they also suggest that the ADA had larger disemployment effects on the disabled in states with more ADA-related charges.

VI. Conclusion

Some social critics see the ADA as part of a process eroding the traditional employment-at-will doctrine and making the U.S. labor market more like labor markets in Europe. In contrast, ADA proponents see the ADA as making the labor market more inclusive, without appreciably increasing employer costs or reducing employment. Economic theory is useful for evaluating these alternative views and suggests avenues for inquiry, but does not make unambiguous predictions.

In 1993, the first full year in which the ADA was in effect, there were marked drops in the employment of disabled men aged 21-58 and disabled women aged 21-39. Extrapolating employment trends, composition effects, and changes in DI and SSI participation rates do not seem to account for this decline, leaving the ADA as a likely cause. This interpretation is also supported by evidence that employment of disabled men fell more sharply in states with more ADA-related charge activity and by relative declines in disabled employment in medium size firms.

Since the ADA provides a form of employment protection, it can lead to lower separation and accession rates for the disabled. There is little evidence of an effect of the ADA on disabled separations, however. This and the fact that costs of reasonable accommodation are probably larger than the costs of litigation for wrongful termination suggests that accommodation costs have been more important for employers than lawsuits. The absence of an offsetting decline in wages suggests the equal pay provision has also played a role in the ADA's employment effects. Finally, we found no evidence of effects on nondisabled workers. Contrary to the concerns of its fiercest opponents (e.g. Olson, 1997), it seems likely that the ADA predominantly affected the costs of employing the disabled, and has not led to a fear-of-litigation climate that reduced the overall level of employment.

Appendix A: Data Issues

1. March 1993 to March 1994 Match

In principle, households in rotation groups 1-4 in 1993 are interviewed in March 1994 when they are in rotation groups 5-8. In practice, some of these households move or are lost for other reasons. For the purposes of the 93/94 match, we selected individuals in the relevant rotation groups, with valid interview status (CPS item H-HHTYPE=1), and in the age range of interest (21-58 for 1993 and 21-60 for 1994). Records were matched using the household identifier (CPS item H-IDNUM), person line number (CPS item A-LINENO), and rotation group (i.e., 1994 H-MIS equaled 1993 H-MIS plus four). Of the March 1993 records eligible for matching, 76.2 percent were matched to a March 1994 record.

We defined a successful match as a person with the same sex and race in both years. 70.9 percent of eligible March 1993 records were matched according to this definition, a rate similar to that reported in Peracchi and Welch (1995). 93 percent of the records that satisfied the basic match restrictions were also matched on sex and race. Because the household identifier on the 1994 ICPSR file is incorrect, the matched data use a corrected household identifier for 1994 provided by the Census Bureau.

2. Consequences of the CPS Redesign

As noted in the text, the redesign changed the main questionnaire and instituted universal computer-assisted interviewing (see, e.g., Polivka (1996)). The supplement questionnaire was unchanged but these other changes may have affected the nature or likelihood of supplement responses. A parallel survey conducted in 1993 showed few differences between computer-assisted and paper-and-pencil interview results for the supplement. Later, however, it was discovered that some of the annual earnings data collected in the 1994 supplement were mistakenly collected for subannual amounts (Bureau of the Census, 1994). The problem appears to have been fixed in later surveys. To minimize the consequences of errors in annual earnings, we exclude observations on weekly wages below \$25 or above \$2000 (in 1988 dollars).

One possible consequence of the redesign was fluctuation in supplement non-interview rates, which were low in the 1994 CPS and high in the 1995 CPS (personal communication from G. Weyland, Bureau of the Census). This affects the data because Supplement variables for people who do not respond are "allocated" by the Census Bureau. Another possible consequence of the redesign is a change in the type of people who report themselves as disabled. To assess the impact of these phenomena, we used the March 1993-94 matched data. The March 1993 data were collected using the old CPS methodology. This allows us to restrict the sample for these two years to people who responded to the CPS supplement in 1993 (i.e., their responses were not allocated), and/or who identified themselves as disabled in both years, and/or who answered the supplement questions on weeks worked in both years.

The results of this analysis are reported in Table A1. Again, the coefficient of interest is a 1994 dummy interacted with a dummy for being disabled. Except for men aged 21-39, results under a variety of sample restrictions are remarkably consistent with the main set of cross-sectional results, though less precisely estimated. The results for men aged 21-39 fluctuate considerably across columns, and in some cases are much larger or smaller than the cross-sectional results in column (1). This variability may be due to the fact that match rates are lower for this age group. Controlling for schooling, marital status, and veteran status reduces the fluctuation in estimates across samples.

3. March to April Matches for 1988-1997

March households in rotation groups 1-3 and 5-7 are eligible for interviewing in April. Observations to be matched were selected on the basis of rotation group, interview status, and age, as described for the March 1993-94 match. Revised Census Bureau confidentiality rules necessitated additional criteria for matching records in some years. We therefore matched households using the household identifier (H-IDNUM), person line number (A-LINENO), rotation group status (i.e., April H-MIS equaled March H-MIS plus one), and 1960 Census state code (HG-ST60). Of the eligible March records, 92.8 percent were successfully matched to an April record using these basic criteria (H-IDNUM, A-LINENO, H-MIS, and HG-ST60); 88.6 percent were successfully matched on both the basic criterion and sex and race, which is 95.4 percent of those matched using only basic criteria.

Appendix B: Effects of SSI and DI on Employment of the Disabled

The CPS identifies individuals receiving SSI and social security income (OASDI). For workers below retirement age, the latter consists of disability benefits and survivor benefits. We coded everyone receiving more than $75 (in 1988 dollars) a week in social security income as receiving DI since survivors benefits are typically smaller than DI for pre-retirement age groups (Department of Health and Human Services, 1992, Tables 5A.5-6). Figure B1 plots the fraction of men and women of different age groups receiving SSI and/or DI. Although increases in the DI and SSI rolls predate the ADA, the timing is close enough that disemployment effects of these programs might still provide an alternative explanation for the decline in disabled employment. The investigation of this possibility is complicated by the fact that disability income is both a cause and a consequence of employment status.

We use the following simultaneous equation model to describe causal relationships between Social Security benefits and employment:

$$y_{it} = x'_{it}\beta_0 + \alpha s_{it} + \gamma_t a_{it} + \varepsilon_{it} \tag{8a}$$

$$s_{it} = x'_{it}\beta_1 + \delta_1 y_{it} + \eta_{it}, \tag{8b}$$

where y_{it} is weeks worked for individual i at time t, x_{it} is a vector of controls, time dummies, and disability status; s_{it} is a dummy indicating federal transfer payments, and a_{it} is the interaction of disability status and a dummy for post-ADA years. The parameters of interest are the interaction coefficients γ_t as before. ε_{it} and η_{it} are error terms that capture unobserved effects that are orthogonal to x_{it} and a_{it}. But ε_{it} and η_{it} are almost certainly negatively correlated, since an individual who works more will generally be less likely to receive disability income even after controlling for observable characteristics. Since DI and SSI are means-tested, we also have $\delta_1 < 0$.

The parameter α captures the possible labor supply consequences of any transfers and is presumably negative. The regressions reported in Table II amount to omitting s_{it} from equation (8a). Figure B1 shows that the proportion receiving SSI and DI increased over the sample, and in fact this increase is greater for those reporting a work-related disability. The estimates of γ_t in Table II may therefore be biased downwards (i.e., too negative).

One strategy to correct this omitted variable bias is to add s_{it} to the regressions reported in Table II, ignoring the fact that benefit receipt is endogenous. We do this in Table B1 by adding a dummy for income from SSI or DI to our basic specification for men and women aged 21-39.[10] The resulting coefficient estimate is large and negative. As a result, the impact of the ADA on men aged 21-39 is reduced, but is still statistically significant. For men aged 40-58, SSI or DI receipt has an even larger impact on weeks worked, so the effects of the ADA are smaller, but still negative and precisely estimated in 1993 and 1994. However, the disemployment effects for this group disappear in later years. As before, the effects for women aged 40-58 appear to start too early.

Next, we expand the model to recognize that federal benefits can have both income and implicit tax rate effects. Instead of the constant coefficient α, the effect of benefit receipt in

[10]These models also include the regressors used to predict wages in other specifications described below. The regressors are age, age squared, a dummy for age 30-39 or age 40-49, two race dummies, dummies for high school graduates, some college and college graduates, dummies for married, separated, widowed, and having served in Vietnam or other military service, state dummies and finally a dummy for answering yes to the question "Did you ever retire or leave a job for health reasons". Adding these covariates to the regressions in Table II does not change the results since the post-ADA/disability interactions are essentially uncorrelated with individual characteristics.

equation (8a) is now assumed to be:

$$\alpha_{it} = \alpha_0 + \alpha_1 \tau_1 \hat{w}_{it},$$

where $\hat{w}_{it}$ is the predicted wage of the individual (as a function of the observables x_{it}), and τ_1 is the tax rate induced by DI or SSI means-testing. α_0 captures the effect of additional unearned income, and $\alpha_1 \tau_1$ captures the fact that after-tax wages are $(1 - \tau_1 s_{it})\hat{w}_{it}$. Both α_0 and $\alpha_1 \tau_1$ are presumably negative. Columns (2) and (8) in Table B1 report the results of estimating the model with tax effects:

$$y_{it} = x'_{it}\beta_0 + \alpha_0 s_{it} + \alpha_1 \tau_1 \hat{w}_{it} s_{it} + \gamma a_{it} + \varepsilon_{it}.$$

Both s_{it} and $\hat{w}_{it} s_{it}$ have negative coefficients. For example, in column (2), for men aged 21-39, the income effect is 13 weeks, and the coefficient on the interaction term suggests that for a worker earning \$500 a week, an increase in the tax rate from 50 percent to 60 percent reduces employment by one week. Estimates of the impact of the ADA in these models are similar to those estimated in regressions without $\hat{w}_{it} s_{it}$.

The estimates so far do not account for the likelihood that benefit receipt is endogenously determined. It is straightforward to show that the resulting estimates of γ_t are biased upwards (i.e., too small in absolute value). We therefore experimented with an instrumental variables strategy. This approach begins by assuming that s_{it} has two components, means-tested and non-means-tested benefits. In particular, let $s_{it} = \hat{s}_{it} + \tilde{s}_{it}$, where $\hat{s}_{it}$ is a means-tested benefit like SSI or DI. Only means-tested benefits should have a tax-rate effect so the estimating equation becomes

$$y_{it} = x'_{it}\beta_0 + \alpha_0 s_{it} + \alpha_1 \tau_1 \hat{w}_{it} \hat{s}_{it} + \gamma a_{it} + \varepsilon_{it}. \tag{9a'}$$

To implement this strategy we set s_{it} equal to an indicator for *any federal stipend*, including veterans benefits, SSI, any OASDI payment, or another federal pension or disability-related payment (military, railroad, and federal employee disability pensions). $\hat{s}_{it}$ indicates any federal stipend that is subject to means-testing, i.e., SSI, DI and any veterans benefits that are means-tested. Means-testing of veterans benefits can be determined from a CPS question that asks whether respondents are required to fill out an income questionnaire for the Department of Veterans Affairs. The instrument for s_{it} is an indicator for individuals receiving veterans benefits who are *not* subject to means-testing, on the presumption that these benefits are not endogenous.[11] The interaction term, $\hat{w}_{it} \hat{s}_{it}$, is instrumented using the interaction of predicted

[11] Means-tested veterans' benefits include some veterans pension and survivor benefits. Non-means-tested veterans benefits include veterans compensation and schooling benefits (see, e.g., Department of Veterans Affairs, 1990)

wages with an indicator for having had a previous disability that led to job loss (see footnote 9). The reasoning here is that a worker who left a job because of a work-related disability is potentially eligible for DI. The regressions include a main effect for having had a previous work-related disability, so the identification comes from the interaction with $\hat{w}_{it}$.

Columns (3), (4), (9), and (10) report OLS estimates of the models to be estimated by two-stage least squares (2SLS). This generates results similar to those using only an indicator for SSI or DI as a control. 2SLS estimates of models without an interaction term, reported in columns (5) and (11), show much smaller effects of disability benefits on weeks worked, and consequently larger disemployment effects of the ADA than in the OLS models.

Finally, columns (6) and (12) report the results of treating both s_{it} and $\hat{w}_{it}\hat{s}_{it}$ as endogenous. These results show smaller effects of federal disability benefits on young workers, but there are precisely estimated implicit tax effects. The ADA still appears to have reduced the employment of disabled men aged 21-39 by 2.3 weeks in 1993. For young disabled women, the implicit tax effect is considerably larger, but the ADA still has a statistically significant negative effects in 1993 and 1994. For men aged 40-58, the estimates are more precise, and show an approximately 2 week decline over the whole sample period. The results for women aged 40-58 again show little evidence of ADA effects after controlling for endogenous benefit receipt.

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Table I: Descriptive Statistics

	Survey Year									
	Age 21 - 39					Age 40 - 58				
	1988	1990	1992	1994	1996	1988	1990	1992	1994	1996
A. Men										
Age	29.93	30.04	30.16	30.36	30.47	48.08	47.78	47.73	47.77	47.76
White	0.85	0.84	0.84	0.83	0.83	0.87	0.87	0.86	0.86	0.86
High School Graduate	0.39	0.39	0.36	0.35	0.34	0.36	0.35	0.31	0.31	0.31
College Graduate	0.24	0.24	0.22	0.23	0.24	0.27	0.28	0.29	0.30	0.30
Employed	0.86	0.87	0.83	0.84	0.85	0.87	0.87	0.85	0.84	0.85
Weeks Worked	43.9	44.5	42.9	42.8	43.9	44.6	44.8	44.0	43.5	43.9
Weekly Wage	444.5	433.9	418.9	424.0	418.8	629.7	617.6	603.6	612.5	579.5
Has a work-related disability	0.048	0.044	0.048	0.054	0.044	0.092	0.098	0.095	0.101	0.098
Receives SSI or DI payments	0.018	0.015	0.020	0.020	0.020	0.035	0.038	0.040	0.046	0.047
N	22591	22849	22317	20733	17262	15130	15624	16091	15846	14538

Note: Year labels refer to survey years. Data are from the March CPS. Wage data are in 1988 dollars.

Table I (continued)

	Survey Year									
	Age 21 - 39					Age 40 - 58				
	1988	1990	1992	1994	1996	1988	1990	1992	1994	1996
B. Women										
Age	29.94	30.05	30.20	30.38	30.49	48.13	47.87	47.77	47.79	47.83
White	0.83	0.82	0.82	0.81	0.81	0.86	0.85	0.85	0.84	0.84
High School Graduate	0.42	0.40	0.36	0.33	0.31	0.45	0.44	0.39	0.37	0.36
College Graduate	0.21	0.22	0.21	0.22	0.24	0.17	0.20	0.21	0.23	0.24
Employed	0.69	0.69	0.69	0.69	0.70	0.66	0.68	0.69	0.71	0.72
Weeks Worked	33.9	34.4	34.5	34.1	34.9	33.0	34.2	34.9	35.5	36.0
Weekly Wage	294.1	296.9	296.6	302.0	302.6	329.5	340.9	350.3	361.1	365.1
Has a work-related disability	0.038	0.038	0.040	0.047	0.048	0.088	0.092	0.094	0.098	0.107
Receives SSI or DI payments	0.020	0.020	0.020	0.031	0.030	0.037	0.037	0.042	0.047	0.052
N	24899	24949	24408	22756	19143	16110	16863	17143	17364	15676

Note: Year-labels refer to survey years. Data are from the March CPS. Wage data are in 1988 dollars.

Table II: Basic Results

	Dep. Var. Weeks Worked						Dep. Var. Log of Weekly Earnings					
	Men			Women			Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
A. Age 21 - 39												
Disability Main Effect	-24.7	-24.9	-24.3	-16.9	-17.1	-17.2	-0.379	-0.337	-0.315	-0.329	-0.295	-0.291
	(0.25)	(0.61)	(0.46)	(0.39)	(0.94)	(0.70)	(0.016)	(0.039)	(0.030)	(0.019)	(0.046)	(0.035)
Disability*1991	-0.66.	-0.80	-0.22	0.52	0.34	0.17	0.010	0.052	0.074	0.128	0.161	0.164
	(0.55)	(0.78)	(0.67)	(0.84)	(1.19)	(1.02)	(0.036)	(0.051)	(0.044)	(0.040)	(0.057)	(0.049)
Disability*1992	0.40	0.22	1.03	-1.86	-2.12	-2.35	0.042	0.102	0.132	0.050	0.096	0.101
	(0.54)	(0.95)	(0.77)	(0.80)	(1.44)	(1.14)	(0.035)	(0.061)	(0.050)	(0.040)	(0.070)	(0.057)
Disability*1993	-1.83	-2.07		-4.06	-4.40		0.020	0.097		-0.011	0.048	
	(0.53)	(1.14)		(0.80)	(1.73)		(0.036)	(0.074)		(0.042)	(0.085)	
Disability*1994	-1.94	-2.23		-3.84	-4.25		-0.080	0.014		0.012	0.085	
	(0.55)	(1.35)		(0.79)	(2.04)		(0.036)	(0.087)		(0.041)	(0.099)	
Disability*1995	-3.05	-3.39		-2.97	-3.46		-0.043	0.068		0.031	0.117	
	(0.58)	(1.57)		(0.79)	(2.36)		(0.039)	(0.101)		(0.049)	(0.114)	
Disability*1996	-2.79	-3.19		-3.72	-4.28		-0.107	0.021		-0.071	0.029	
	(0.56)	(1.77)		(0.81)	(2.69)		(0.036)	(0.114)		(0.040)	(0.130)	
Disability*1993-1996			-1.31			-4.48			0.101			0.080
			(0.99)			(1.48)			(0.066)			(0.075)
Disability*Linear Trend		0.05	-0.18		0.07	0.14		-0.017	-0.026		-0.013	-0.015
		(0.22)	(0.16)		(0.34)	(0.23)		(0.014)	(0.010)		(0.017)	(0.012)
No. of Observations	193317	193317	193317	211910	211910	211910	167974	167974	167974	157688	157688	157688

Notes: Standard errors in parentheses. The sample includes men or women aged 21-39 in the survey year . The dependent variable in the first six columns is weeks worked. In the last six columns, it is log average weekly earnings. All regressions include year, age and race dummies.

Table II (continued)

	Dep. Var. Weeks Worked						Dep. Var. Log of Weekly Earnings					
	Men			Women			Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
B. Age 40 - 58												
Disability Main Effect	-31.0	-30.4	-31.4	-23.7	-22.4	-21.9	-0.480	-0.399	-0.464	-0.356	-0.443	-0.318
	(0.20)	(0.50)	(0.37)	(0.31)	(0.76)	(0.55)	(0.016)	(0.041)	(0.030)	(0.019)	(0.049)	(0.034)
Disability*1991	0.37	0.90	-0.03	-1.56	-0.38	0.18	-0.047	0.032	-0.031	-0.082	-0.164	-0.046
	(0.44)	(0.62)	(0.53)	(0.66)	(0.94)	(0.79)	(0.035)	(0.051)	(0.043)	(0.042)	(0.060)	(0.050)
Disability*1992	0.91	1.67	0.36	-1.72	-0.05	0.74	-0.136	-0.025	-0.113	-0.002	-0.118	0.049
	(0.42)	(0.75)	(0.59)	(0.67)	(1.15)	(0.91)	(0.034)	(0.062)	(0.049)	(0.042)	(0.074)	(0.057)
Disability*1993	-1.79	-0.81		-0.35	1.80		-0.023	0.120		0.007	-0.143	
	(0.42)	(0.91)		(0.64)	(1.37)		(0.036)	(0.075)		(0.039)	(0.087)	
Disability*1994	-1.25	-0.05		-0.13	2.50		-0.039	0.136		-0.056	-0.239	
	(0.41)	(1.07)		(0.61)	(1.61)		(0.034)	(0.088)		(0.037)	(0.102)	
Disability*1995	-1.37	0.04		-2.99	0.13		-0.018	0.189		-0.107	-0.324	
	(0.41)	(1.24)		(0.60)	(1.86)		(0.034)	(0.101)		(0.037)	(0.119)	
Disability*1996	-0.23	1.41		-2.36	1.24		0.016	0.256		-0.137	-0.387	
	(0.40)	(1.41)		(0.59)	(2.12)		(0.033)	(0.115)		(0.036)	(0.135)	
Disability*1993-1996			-2.10			2.77			0.024			0.012
			(0.77)			(1.16)			(0.064)			(0.071)
Disability*Linear Trend		-0.22	0.16		-0.48	-0.71		-0.032	-0.006		0.034	-0.015
		(0.18)	(0.12)		(0.27)	(0.18)		(0.015)	(0.010)		(0.018)	(0.011)
No. of Observations	146309	146309	146309	157589	157589	157589	116665	116665	116665	109082	109082	109082

Notes: Standard errors in parentheses. The sample includes men or women aged 40-58 in the survey year . The dependent variable in the first four columns is weeks worked. In the last four columns, it is log average weekly earnings. All regressions include year, age and race dummies.

Table III: Post-ADA Interactions with EEOC Charge Rates

	Men Aged 21-39		Women Aged 21-39		Men Aged 40-58	
	no trend	trend	no trend	trend	no trend	trend
Year	(1)	(2)	(3)	(4)	(5)	(6)
			A. Disabled			
1993	-2.62	-2.48	2.60	2.30	-5.87	-6.68
	(1.26)	(1.43)	(1.84)	(2.09)	(.98)	(1.10)
1994	-1.28	-1.10	-.23	-.35	-.41	-.68
	(.32)	(.40)	(.45)	(.57)	(.23)	(.29)
1995	-.33	-.08	-.22	-.34	-.37	-.71
	(.33)	(.43)	(.44)	(.61)	(.21)	(.31)
1996	.45	.73	-1.30	-1.61	-.48	-.93
	(.35)	(.51)	(.49)	(.73)	(.23)	(.36)
			B. Nondisabled			
1993	.55	-.21	.14	-.53	.69	.16
	(.29)	(.33)	(.41)	(.45)	(.32)	(.36)
1994	.03	-.22	.16	-.09	.12	-.06
	(.06)	(.09)	(.10)	(.12)	(.08)	(.09)
1995	.07	-.21	.23	-.04	.19	-.04
	(.07)	(.09)	(.09)	(.13)	(.07)	(.10)
1996	.05	-.31	.14	-.20	-.01	-.31
	(.07)	(.11)	(.10)	(.15)	(.07)	(.11)

Note: Standard errors in parantheses. The table reports coefficients from a regression of weeks worked on the state level ADA charge rate of the previous year (per 1,000 disabled people aged 21-59 in 1992). These models also include full sets of age and race main effects, and full sets of year*disability and state*disability interactions.

Table A1
Robustness of Basic Results to Sample Restrictions

	Full Sample	Matched March 1993-March 1994 Data					
		All Matched Records	Same Disability Status in 1993&1994	Valid Supplement Response in 1993	Valid Supplement Response in 1993&1994	Valid Supplement Response & Same Disability in 1993&1994	Previous Column Excluding Allocated Weeks Worked
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
			Men Aged 21-39				
Without Covariates	-1.23	-3.11	-0.30	-3.47	-3.57	-0.04	-0.87
	(0.98)	(1.10)	(1.35)	(1.13)	(1.18)	(1.41)	(1.43)
With Covariates	-1.54	-3.16	-0.98	-3.42	-3.55	-0.67	-1.27
	(0.95)	(1.05)	(1.31)	(1.09)	(1.13)	(1.37)	(1.38)
N	19271	13251	12669	12072	11247	10811	10656
			Men Aged 40-58				
Without Covariates	-3.75	-3.62	-3.40	-4.18	-3.68	-2.76	-2.65
	(0.77)	(0.83)	(0.91)	(0.86)	(0.88)	(0.93)	(0.93)
With Covariates	-4.15	-3.99	-3.58	-4.52	-4.06	-2.99	-2.94
	(0.74)	(0.79)	(0.89)	(0.82)	(0.85)	(0.91)	(0.91)
N	15105	12719	11914	11488	10637	10060	9951
			Women Aged 21-39				
Without Covariates	-2.69	-1.13	-2.50	-1.75	-1.51	-2.46	-2.54
	(1.39)	(1.69)	(2.28)	(1.75)	(1.81)	(2.33)	(2.35)
With Covariates	-2.29	-1.24	-2.79	-1.90	-1.78	-2.76	-2.85
	(1.35)	(1.63)	(2.21)	(1.70)	(1.75)	(2.26)	(2.27)
N	21372	15130	14467	13878	12958	12469	12388
			Women Aged 40-58				
Without Covariates	1.09	1.55	-0.42	2.32	2.21	-0.17	-0.19
	(1.12)	(1.24)	(1.55)	(1.30)	(1.35)	(1.61)	(1.62)
With Covariates	1.01	1.38	-0.59	2.27	2.18	-0.36	-0.38
	(1.08)	(1.20)	(1.51)	(1.27)	(1.31)	(1.57)	(1.57)
N	16407	13925	12907	12523	11522	10807	10699

Notes : Standard errors are shown in parentheses. All entries are OLS estimates of coefficients on the disabled*1994 dummy in equations for weeks worked. Models without covariates include year, age, race, and disabled main effects. Models with covariates also contain education, marital status, and veteran status controls.

Table B1: The Role of DI and SSI: Dependent Variable Weeks Worked

	Men						Women					
	(1) OLS	(2) OLS	(3) OLS	(4) OLS	(5) IV	(6) IV	(7) OLS	(8) OLS	(9) OLS	(10) OLS	(11) IV	(12) IV
A. Age 21 - 39												
Disability	-15.04	-15.31	-15.60	-15.74	-19.44	-18.20	-10.44	-10.76	-10.54	-10.84	-12.43	-7.48
	(0.27)	(0.27)	(0.27)	(0.27)	(0.30)	(0.72)	(0.39)	(0.39)	(0.39)	(0.39)	(0.54)	(0.95)
Disability*1991	-0.68	-0.67	-1.14	-0.97	-1.19	-1.06	0.96	1.13	0.90	1.11	0.81	2.36
	(0.52)	(0.52)	(0.52)	(0.52)	(0.53)	(0.53)	(0.80)	(0.80)	(0.80)	(0.80)	(0.80)	(0.85)
Disability*1992	0.78	0.78	0.59	0.76	0.28	0.50	-1.79	-1.79	-1.61	-1.64	-1.97	-0.86
	(0.51)	(0.51)	(0.51)	(0.51)	(0.52)	(0.53)	(0.77)	(0.77)	(0.77)	(0.77)	(0.77)	(0.80)
Disability*1993	-1.74	-1.87	-2.00	-2.15	-2.29	-2.28	-3.43	-3.50	-3.39	-3.48	-3.90	-2.61
	(0.51)	(0.51)	(0.51)	(0.51)	(0.51)	(0.51)	(0.76)	(0.76)	(0.76)	(0.76)	(0.77)	(0.81)
Disability*1994	-0.94	-1.08	-1.36	-1.42	-2.05	-1.85	-3.09	-3.09	-3.15	-3.12	-3.69	-1.62
	(0.52)	(0.52)	(0.52)	(0.52)	(0.53)	(0.53)	(0.76)	(0.76)	(0.76)	(0.75)	(0.76)	(0.84)
Disability*1995	-1.30	-0.15	-2.02	0.13	-2.70	-1.03	-1.84	-1.64	-1.73	-1.54	-2.37	1.00
	(0.55)	(0.56)	(0.55)	(0.56)	(0.55)	(1.03)	(0.75)	(0.75)	(0.75)	(0.75)	(0.76)	(0.94)
Disability*1996	-1.74	-1.65	-2.30	-1.96	-3.25	-2.69	-2.20	-1.79	-2.28	-1.86	-3.01	2.03
	(0.53)	(0.53)	(0.53)	(0.53)	(0.54)	(0.61)	(0.78)	(0.78)	(0.78)	(0.78)	(0.79)	(1.14)
SSI or DI	-17.71	-13.00					-13.91	-7.69				
	(0.27)	(0.55)					(0.32)	(0.88)				
SSI or DI*predicted wage		-0.02						-0.03				
		(0.002)						(0.004)				
Any Federal Stipend			-12.87	-6.93	-1.35	-1.40			-12.22	-7.03	-4.87	-0.60
			(0.22)	(0.33)	(0.47)	(0.47)			(0.27)	(0.69)	(1.44)	(1.65)
Means-tested Fed. Stipend*pred. wage				-0.03		-0.02				-0.02		-0.14
				(0.001)		(0.01)				(0.003)		(0.02)

Note: Standard errors in parentheses. SSIorDI is a dummy which takes the value 1 if the individual receives SSI income or DI income more than $75 a week (in 1988 dollars). Any Fed.Stipend is a dummy for when the individual receives any federal disability income. Pred. Wage is the predicted wage of the individual (see text).

Table B1 (continued)

	Men						Women					
	(1) OLS	(2) OLS	(3) OLS	(4) OLS	(5) IV	(6) IV	(7) OLS	(8) OLS	(9) OLS	(10) OLS	(11) IV	(12) IV
B. Age 40 - 58												
Disability	-17.71	-17.84	-17.82	-17.79	-21.65	-22.73	-16.02	-16.31	-15.44	-15.76	-17.13	-13.26
	(0.22)	(0.22)	(0.22)	(0.22)	(0.24)	(0.35)	(0.32)	(0.32)	(0.32)	(0.32)	(0.45)	(0.61)
Disability*1991	0.58	.0.38	0.09	-0.11	0.18	0.31	-1.06	-0.80	-1.13	-0.85	-1.34	0.53
	(0.41)	(0.41)	(0.41)	(0.41)	(0.42)	(0.43)	(0.63)	(0.63)	(0.63)	(0.63)	(0.63)	(0.67)
Disability*1992	1.87	1.45	1.16	0.51	0.89	1.16	-1.24	-1.28	-1.48	-1.45	-1.66	-0.95
	(0.39)	(0.39)	(0.40)	(0.39)	(0.40)	(0.42)	(0.64)	(0.64)	(0.63)	(0.63)	(0.64)	(0.65)
Disability*1993	-1.03	-1.15	-1.88	-1.90	-2.22	-2.30	0.15	0.56	0.08	0.48	-0.24	2.49
	(0.39)	(0.39)	(0.40)	(0.39)	(0.40)	(0.41)	(0.61)	(0.61)	(0.61)	(0.61)	(0.61)	(0.68)
Disability*1994	-1.32	-1.23	-2.05	-1.56	-2.26	-2.57	-0.40	-0.39	-0.66	-0.62	-0.75	-0.29
	(0.38)	(0.38)	(0.39)	(0.38)	(0.39)	(0.41)	(0.59)	(0.59)	(0.59)	(0.59)	(0.59)	(0.60)
Disability*1995	-0.62	-0.36	-1.31	0.54	-1.89	-2.46	-1.97	-1.71	-2.06	-1.79	-2.51	0.06
	(0.38)	(0.38)	(0.38)	(0.38)	(0.39)	(0.42)	(0.57)	(0.57)	(0.57)	(0.57)	(0.58)	(0.64)
Disability*1996	0.60	1.00	-0.56	0.60	-1.11	-1.88	-1.42	-1.78	-1.73	-1.97	-2.31	-1.53
	(0.38)	(0.38)	(0.38)	(0.38)	(0.39)	(0.43)	(0.57)	(0.57)	(0.56)	(0.57)	(0.57)	(0.59)
SSI or DI	-19.12	-14.60					-15.36	-8.48				
	(0.20)	(0.37)					(0.28)	(0.67)				
SSI or DI*predicted wage		-0.01						-0.03				
		(0.001)						(0.003)				
Any Federal Stipend			-12.70	-6.65	-0.89	-0.80			-14.95	-9.80	-8.20	-6.80
			(0.16)	(0.21)	(0.29)	(0.30)			(0.24)	(0.51)	(1.28)	(1.29)
Means-tested Fed. Stipend*pred. wage				-0.03		0.01				-0.02		-0.10
				(0.001)		(0.003)				(0.002)		(0.009)

Note: Standard errors in parentheses. SSIorDI is a dummy which takes the value 1 if the individual receives SSI income or DI income more than $75 a week (in 1988 dollars). Any Fed.Stipend is a dummy for when the individual receives any federal disability income. Pred. Wage is the predicted wage of the individual (see text).

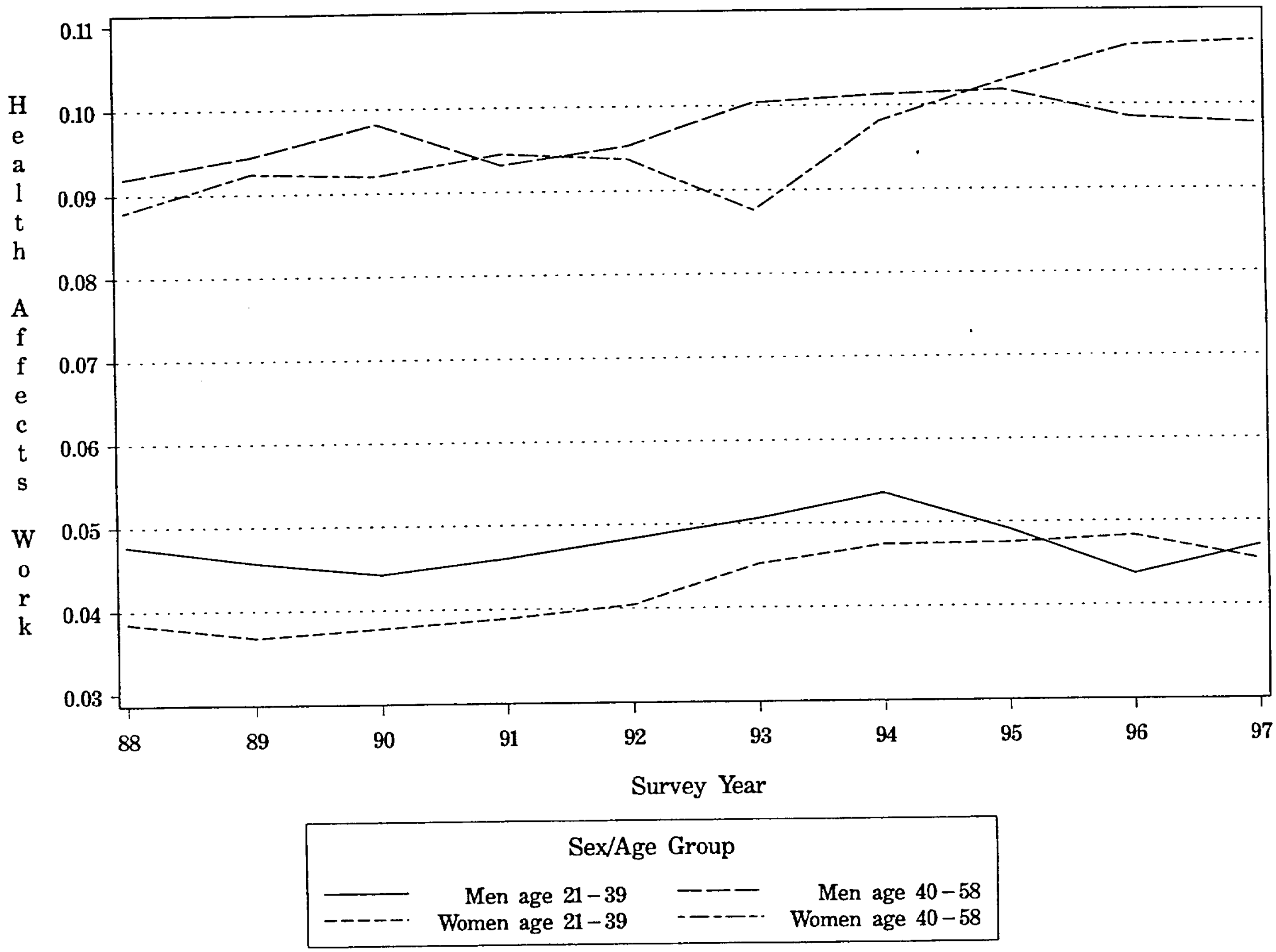

Figure 1. Work – related disability rates for men and women aged 21 – 58 in the 1988 – 97 March CPS.

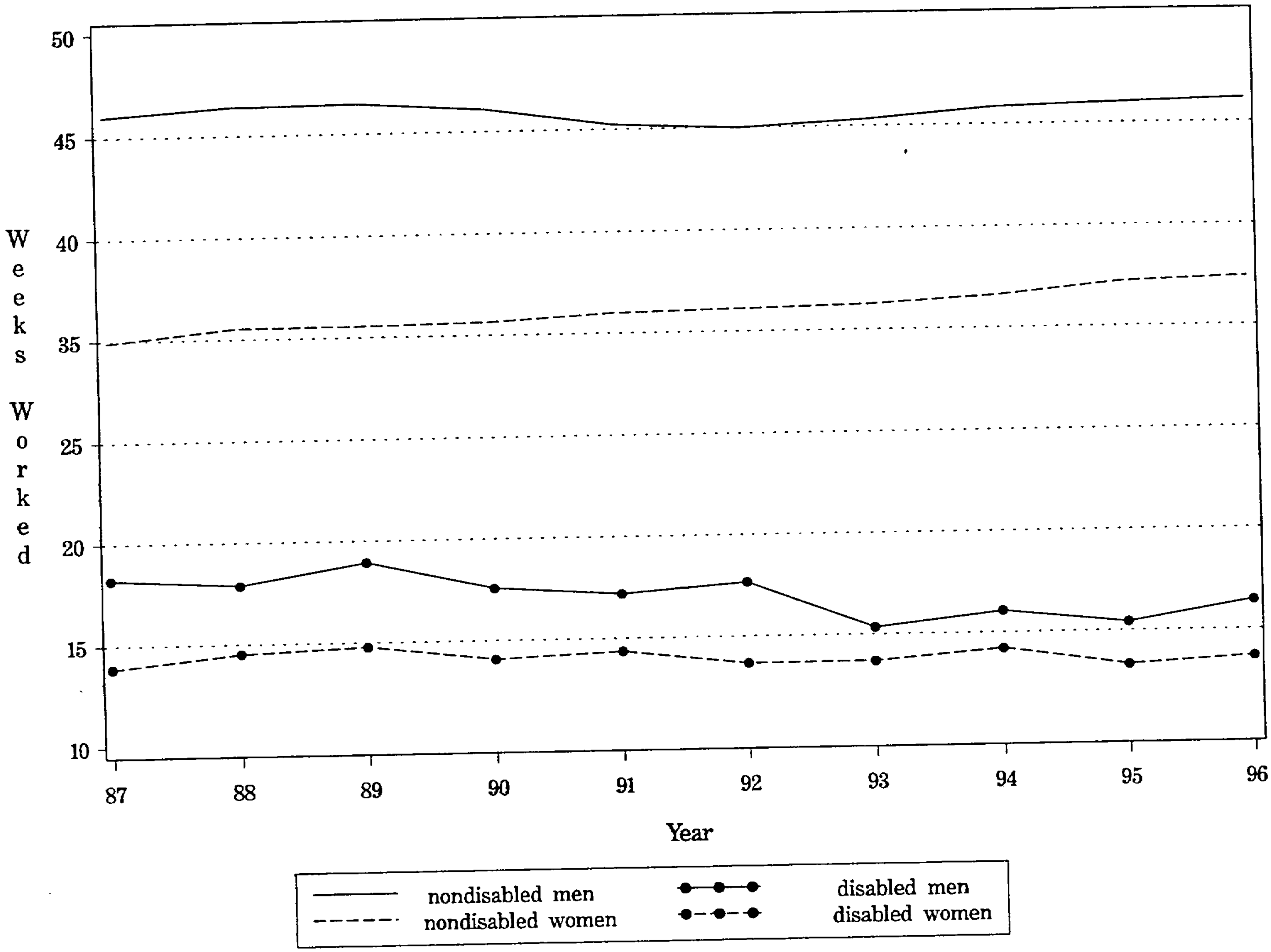

Figure 2. Weeks worked last year by sex and disability status, CPS respondents aged 21–58.

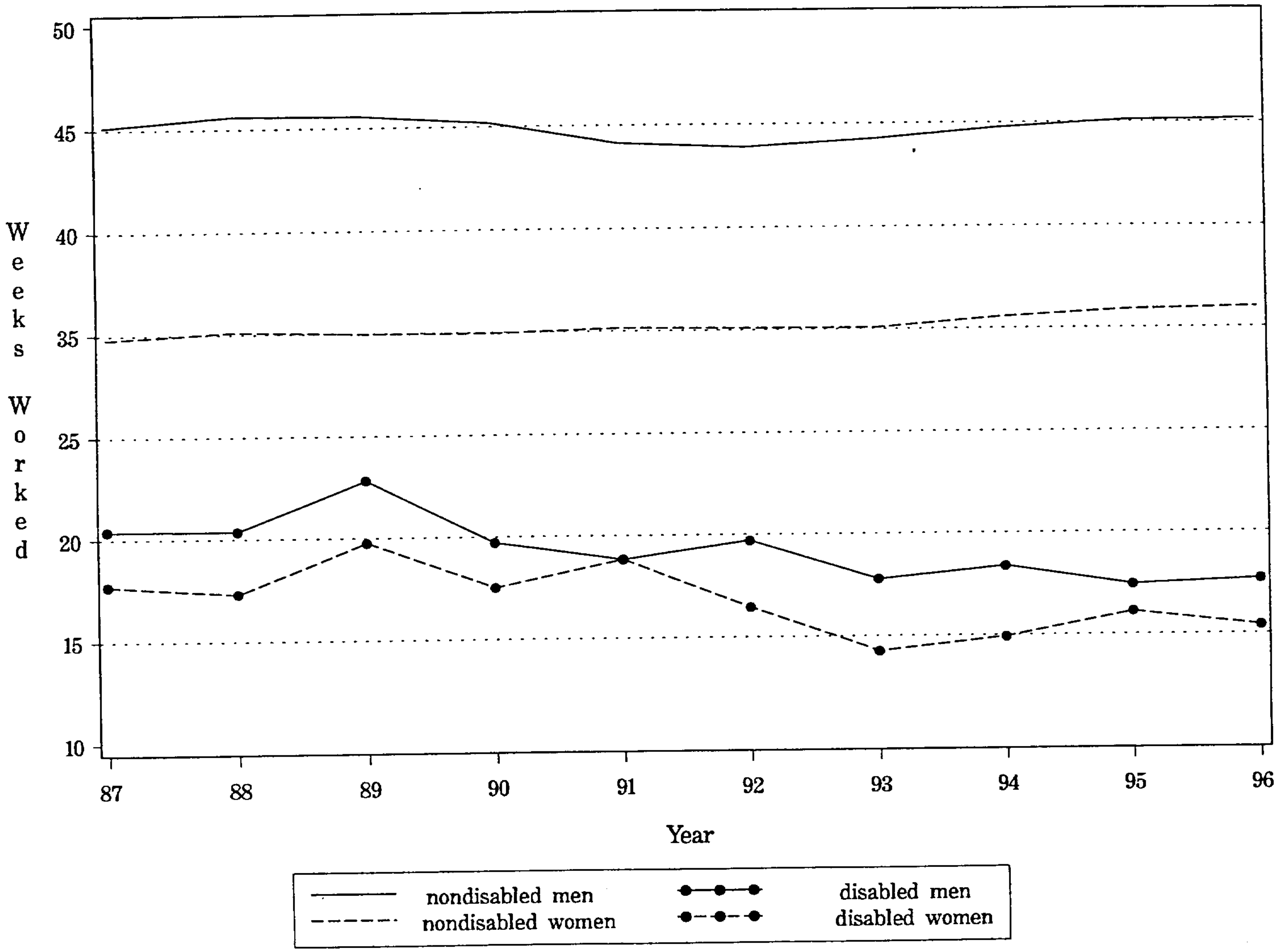

Figure 3a. Weeks worked last year by sex and disability status, CPS respondents aged 21–39.

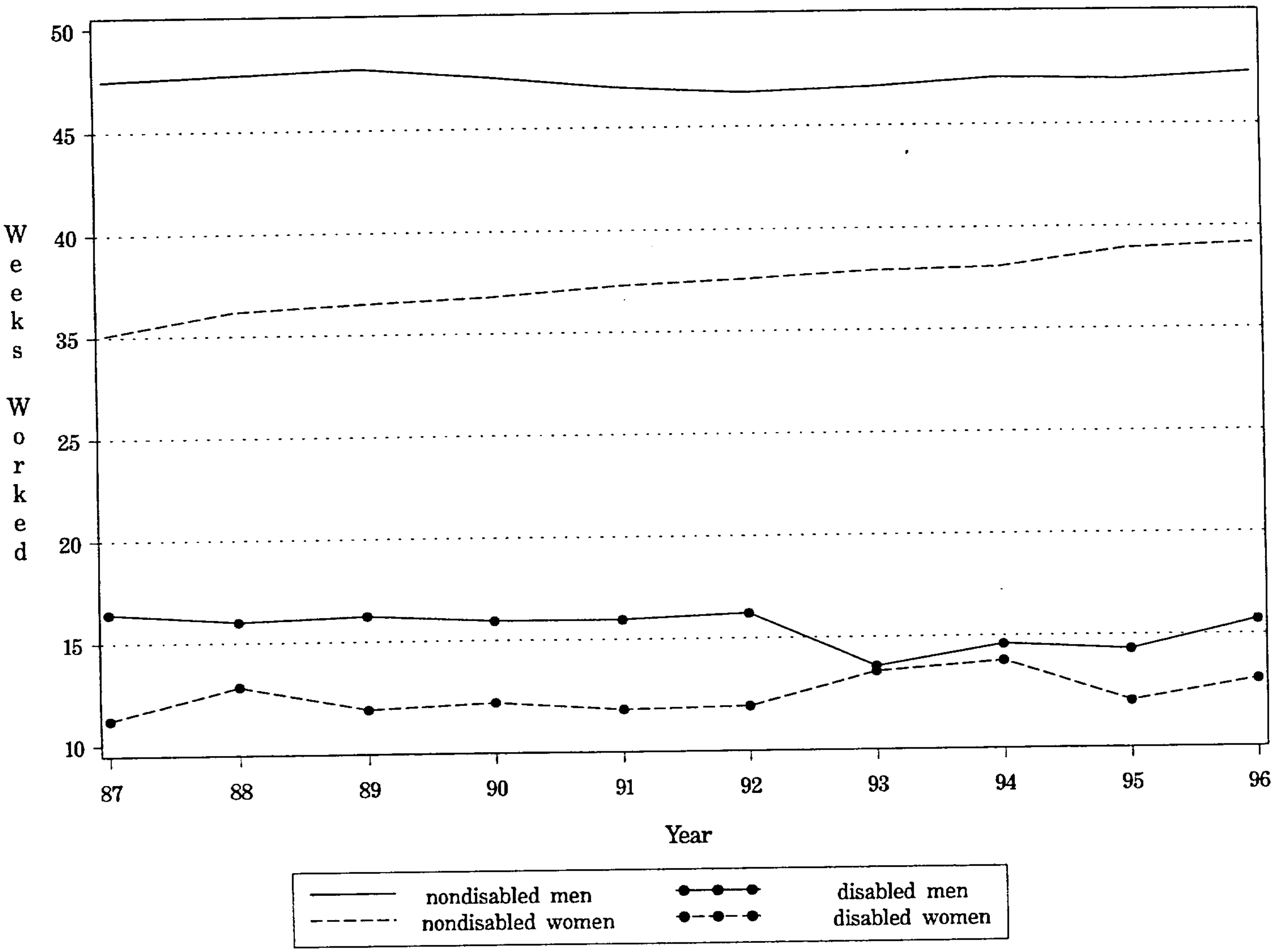

Figure 3b. Weeks worked last year by sex and disability status, CPS respondents aged 40–58.

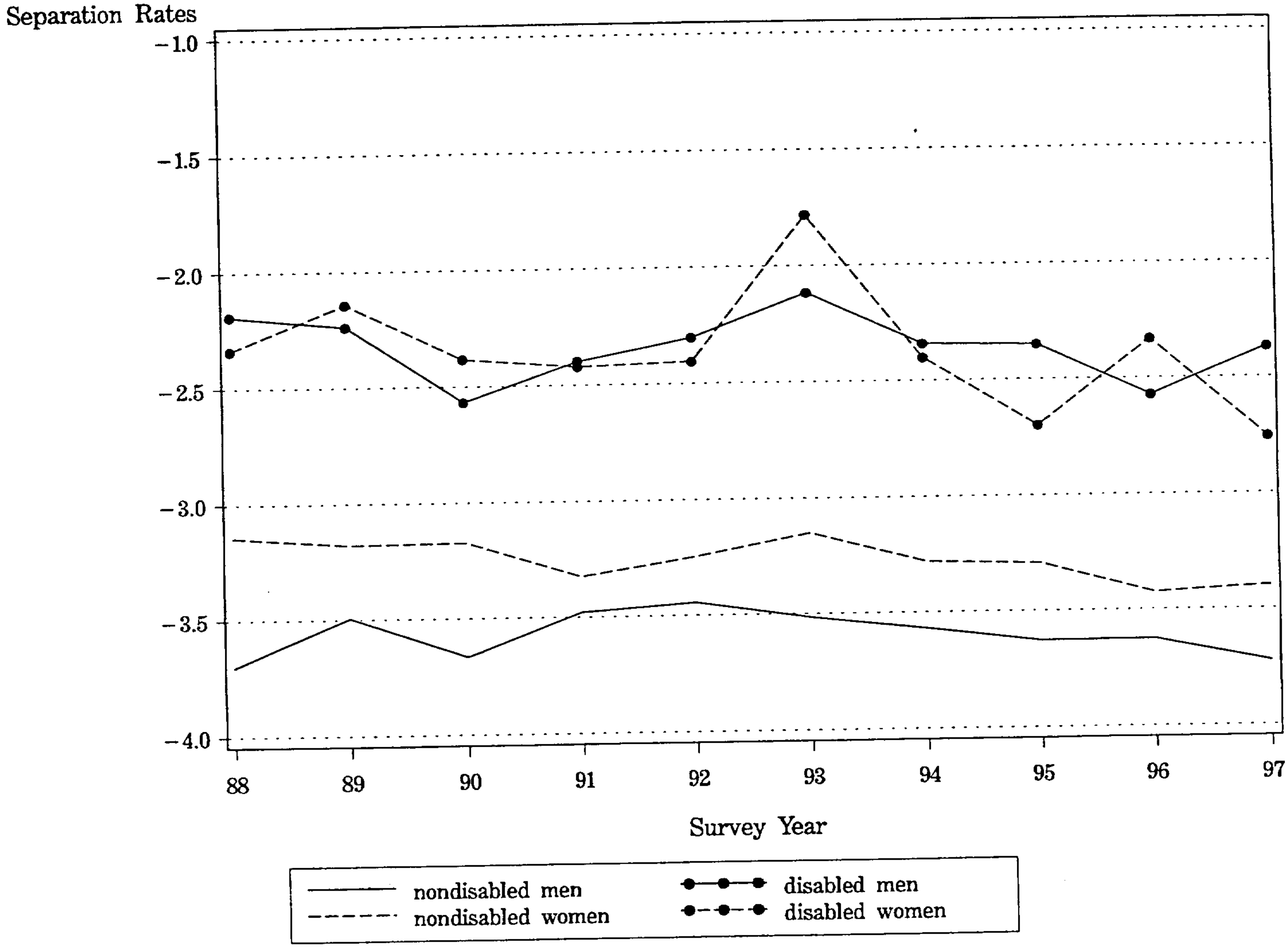

Figure 4a. (Log) Separation rates by disability status and sex. CPS respondents aged 21−39.

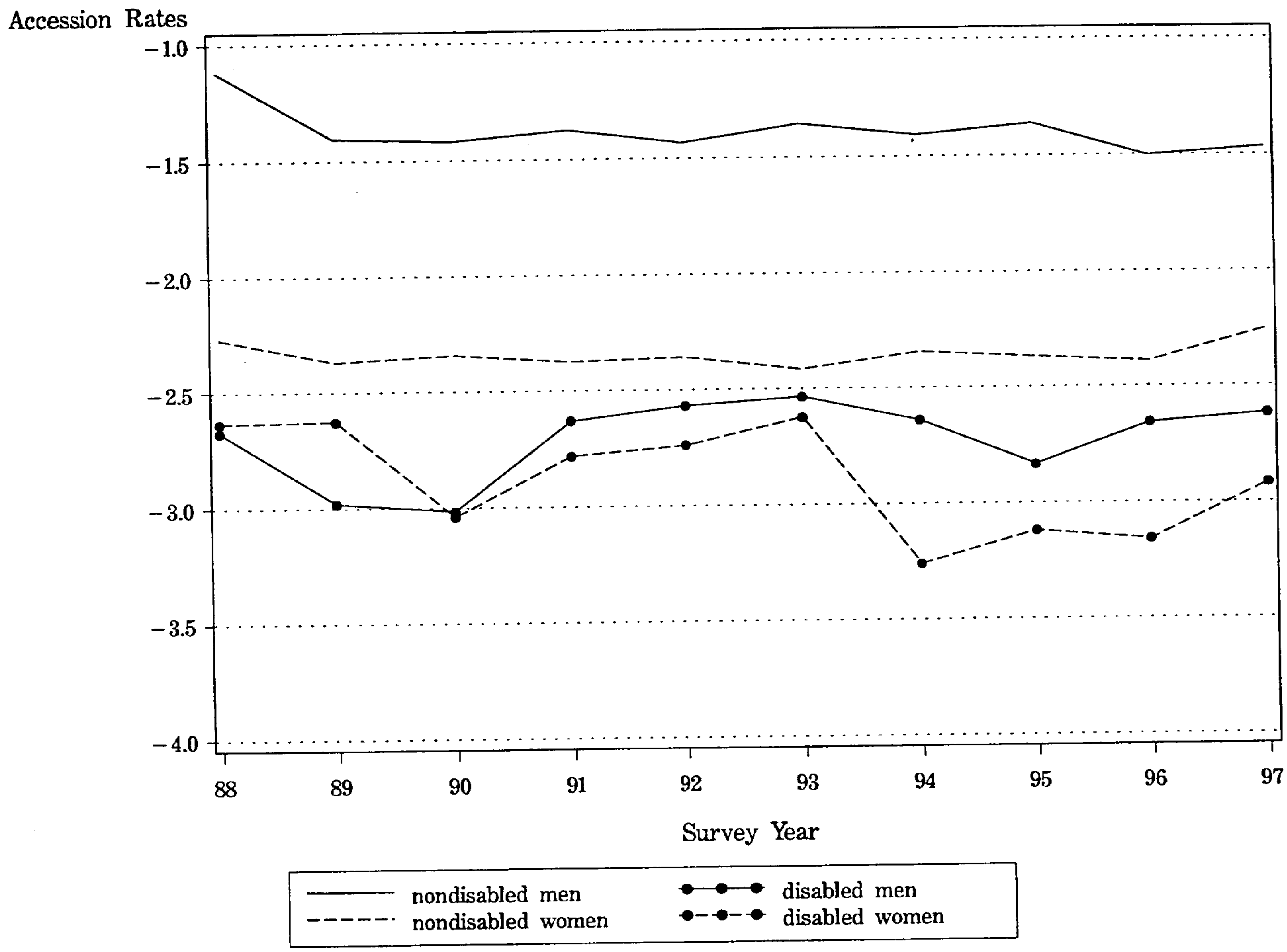

Figure 4b. (Log) Accession rates by disability status and sex. CPS respondents aged 21−39.

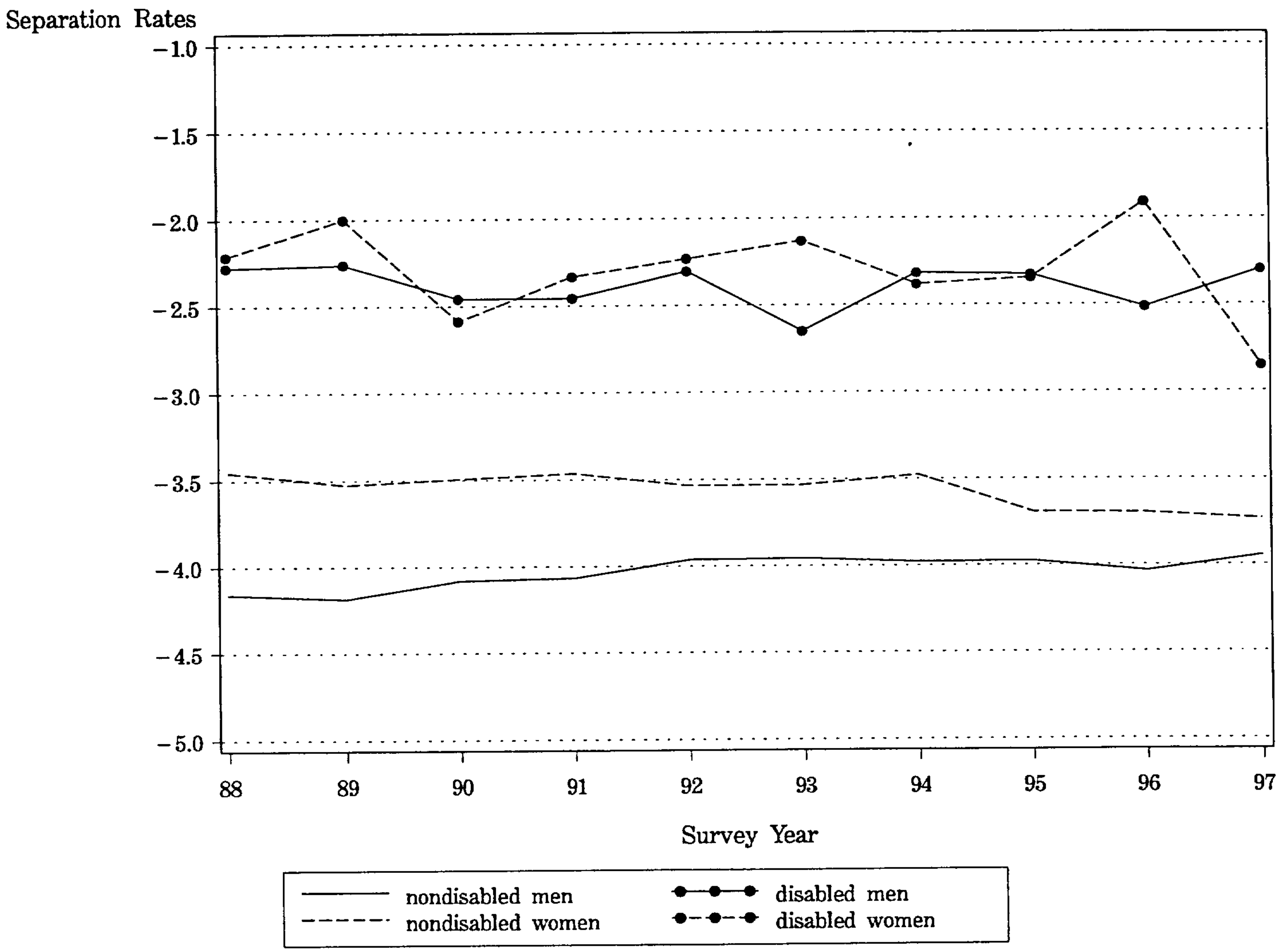

Figure 4c. (Log) Separation rates by disability status and sex. CPS respondents aged 40−58.

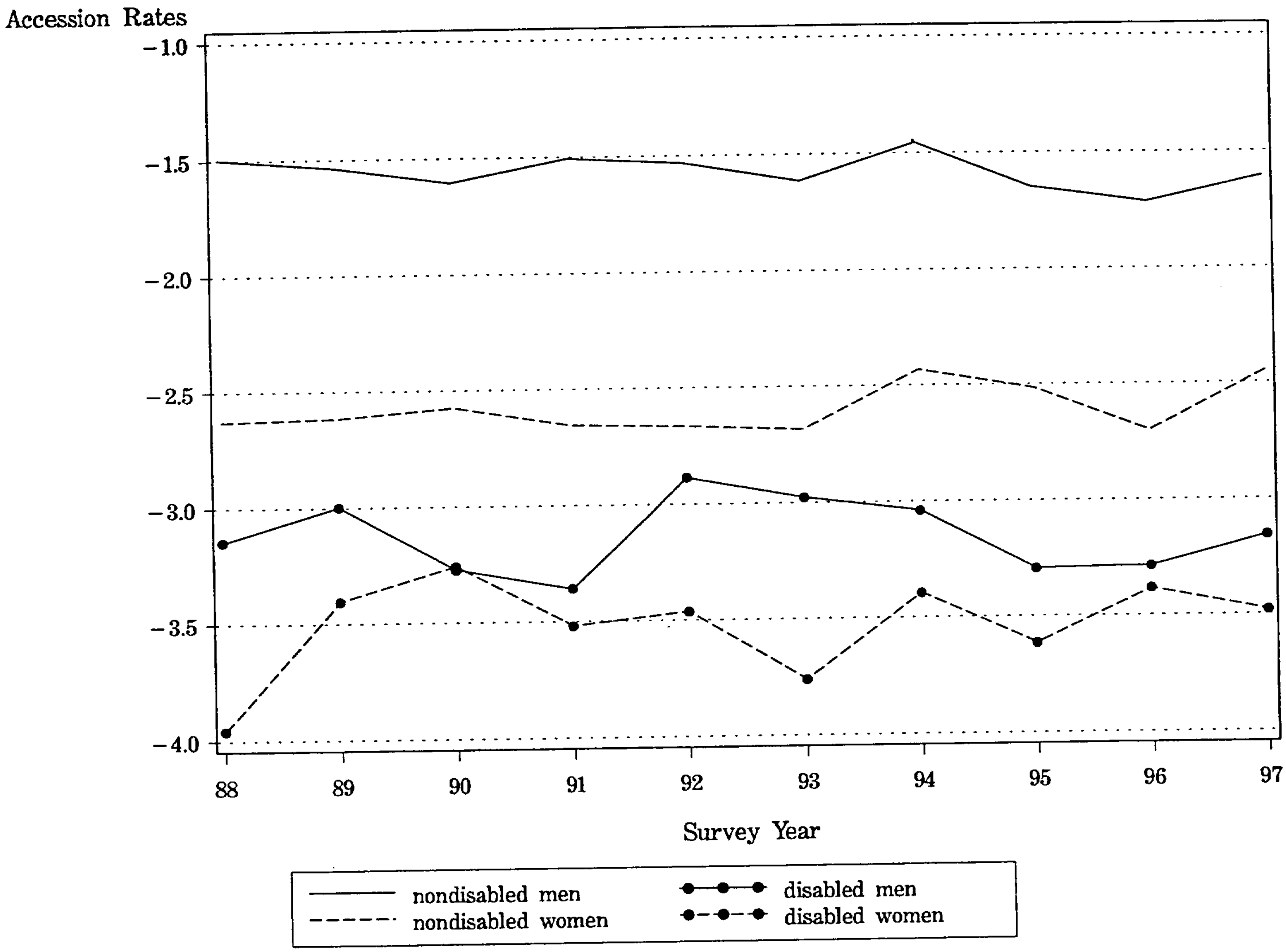

Figure 4d. (Log) Accession rates by disability status and sex. CPS respondents aged 40 – 58.

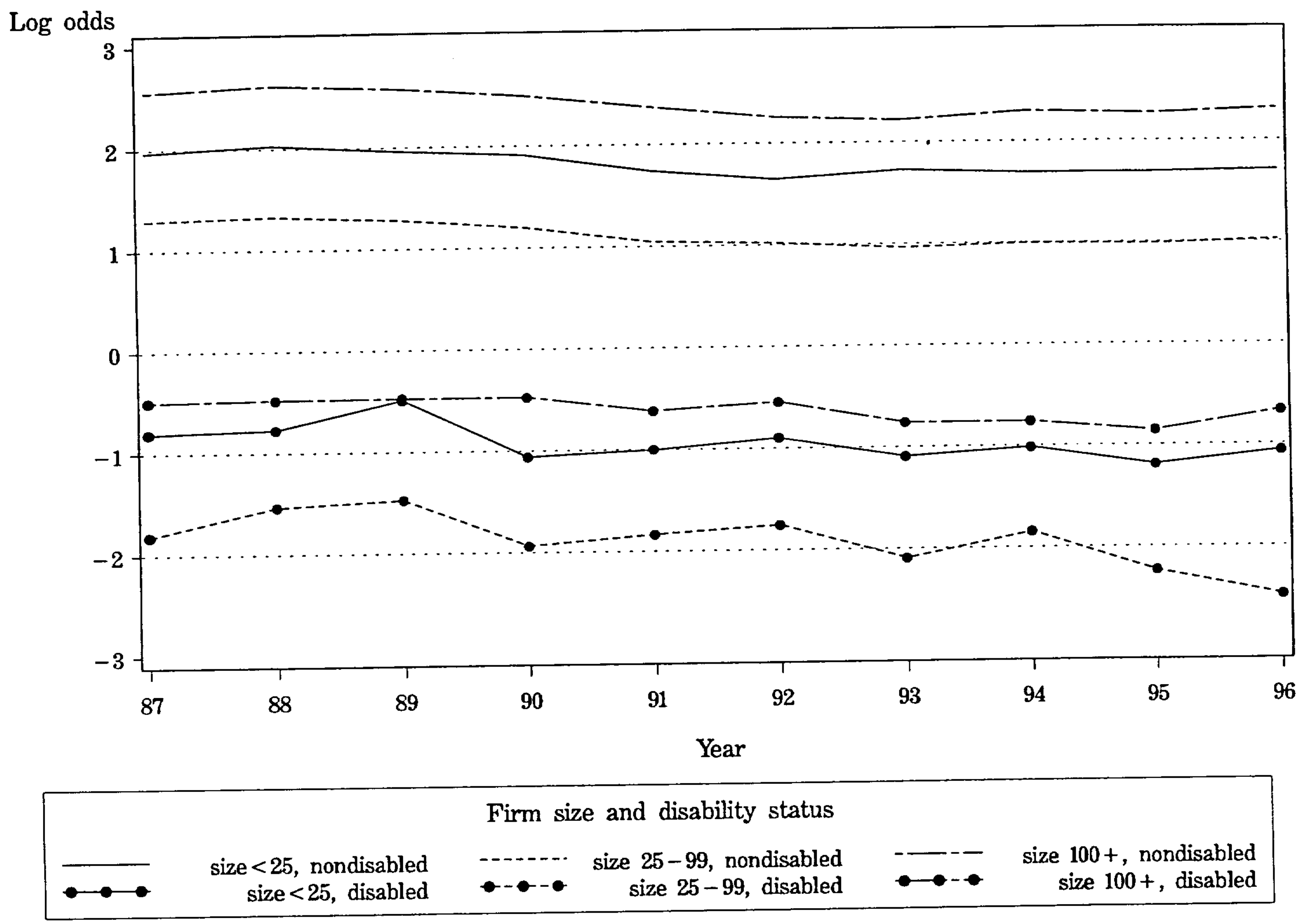

Figure 5a. MNL plot of log employment probabilities by firm size and disability status for men aged 21 − 39. Reference group consists of nonworkers.

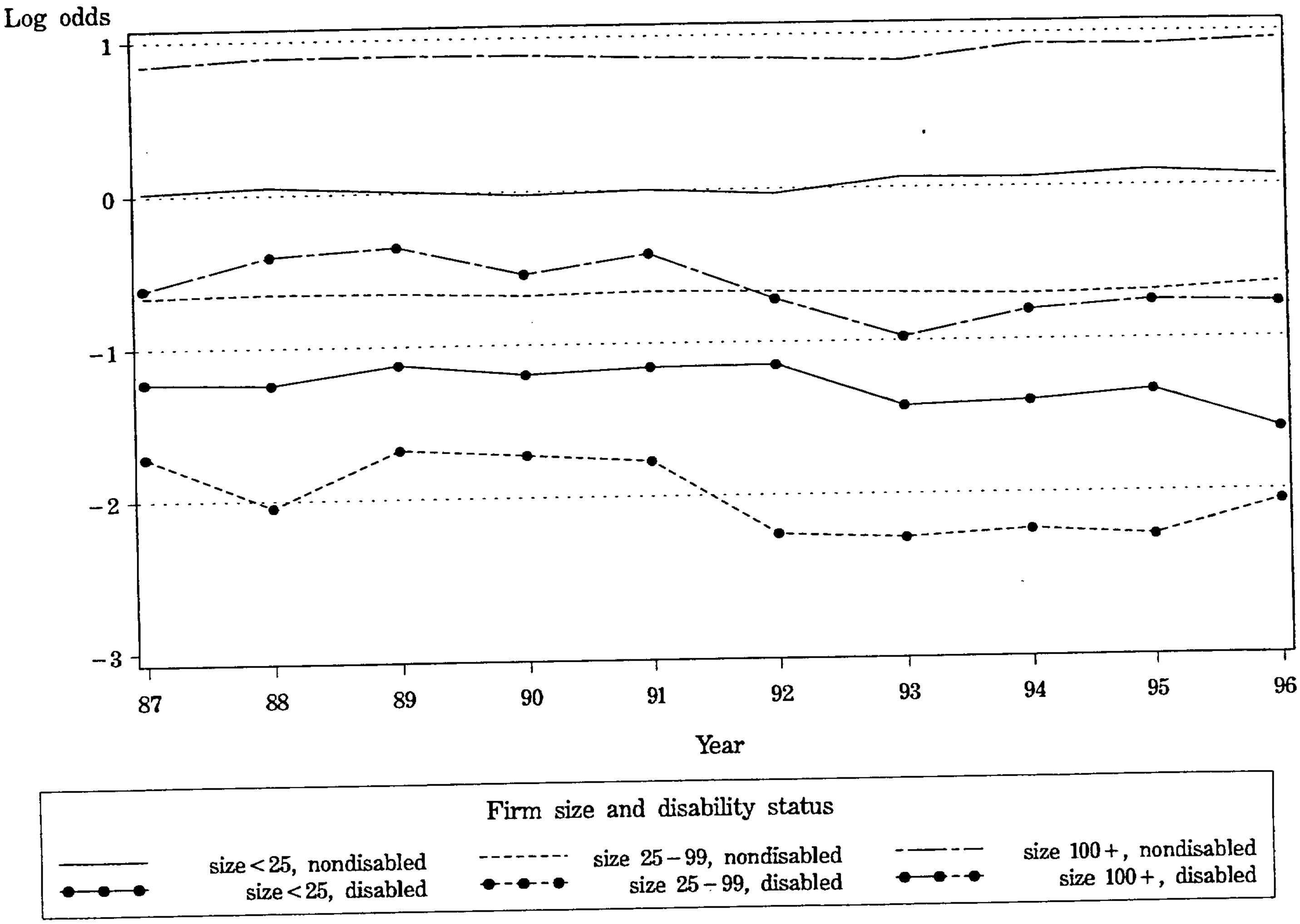

Figure 5b. MNL plot of log employment probabilities by firm size and disability status for women aged 21 − 39. Reference group consists of nonworkers.

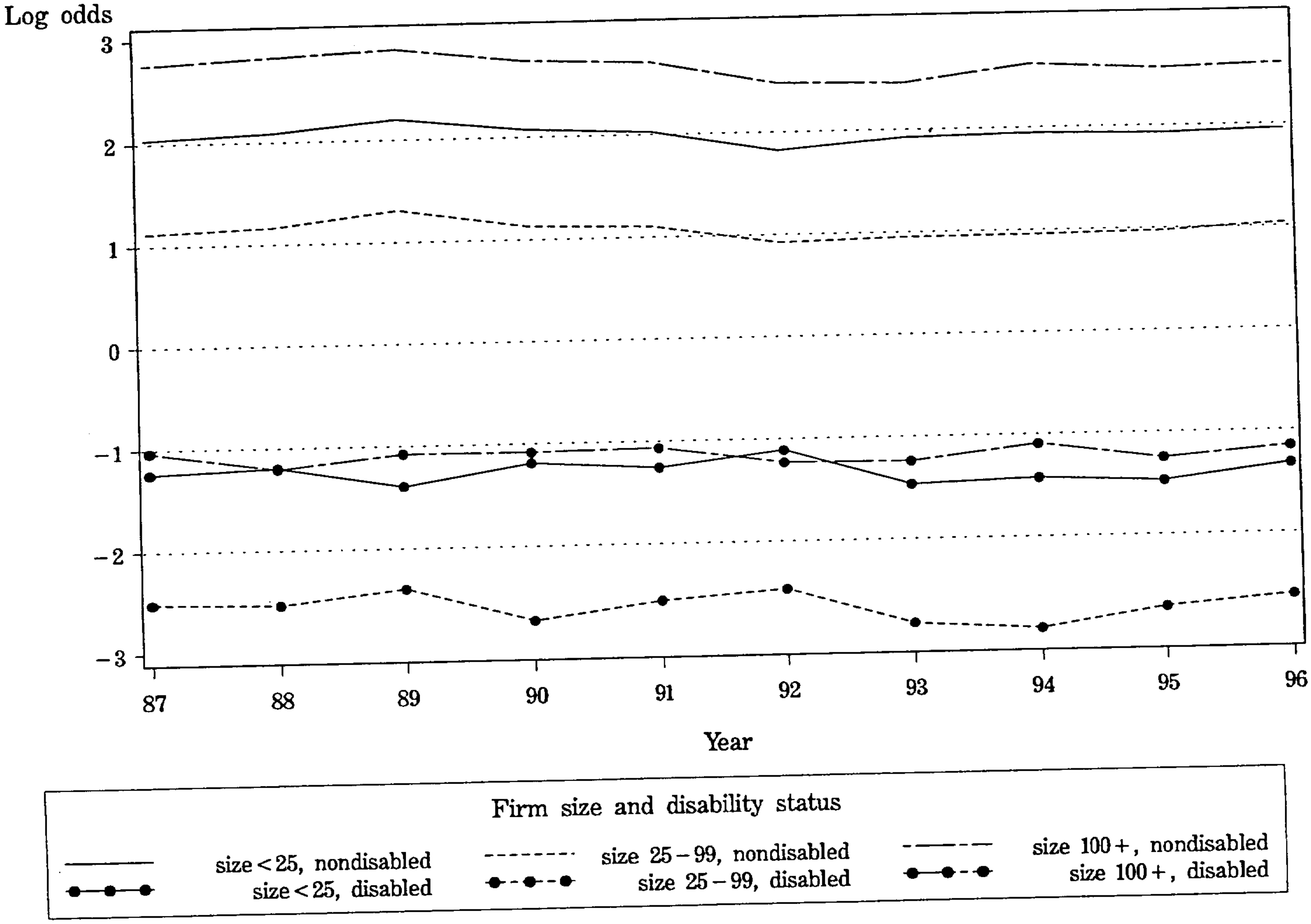

Figure 5c. MNL plot of log employment probabilities by firm size and disability status for men aged 40 − 58. Reference group consists of nonworkers.

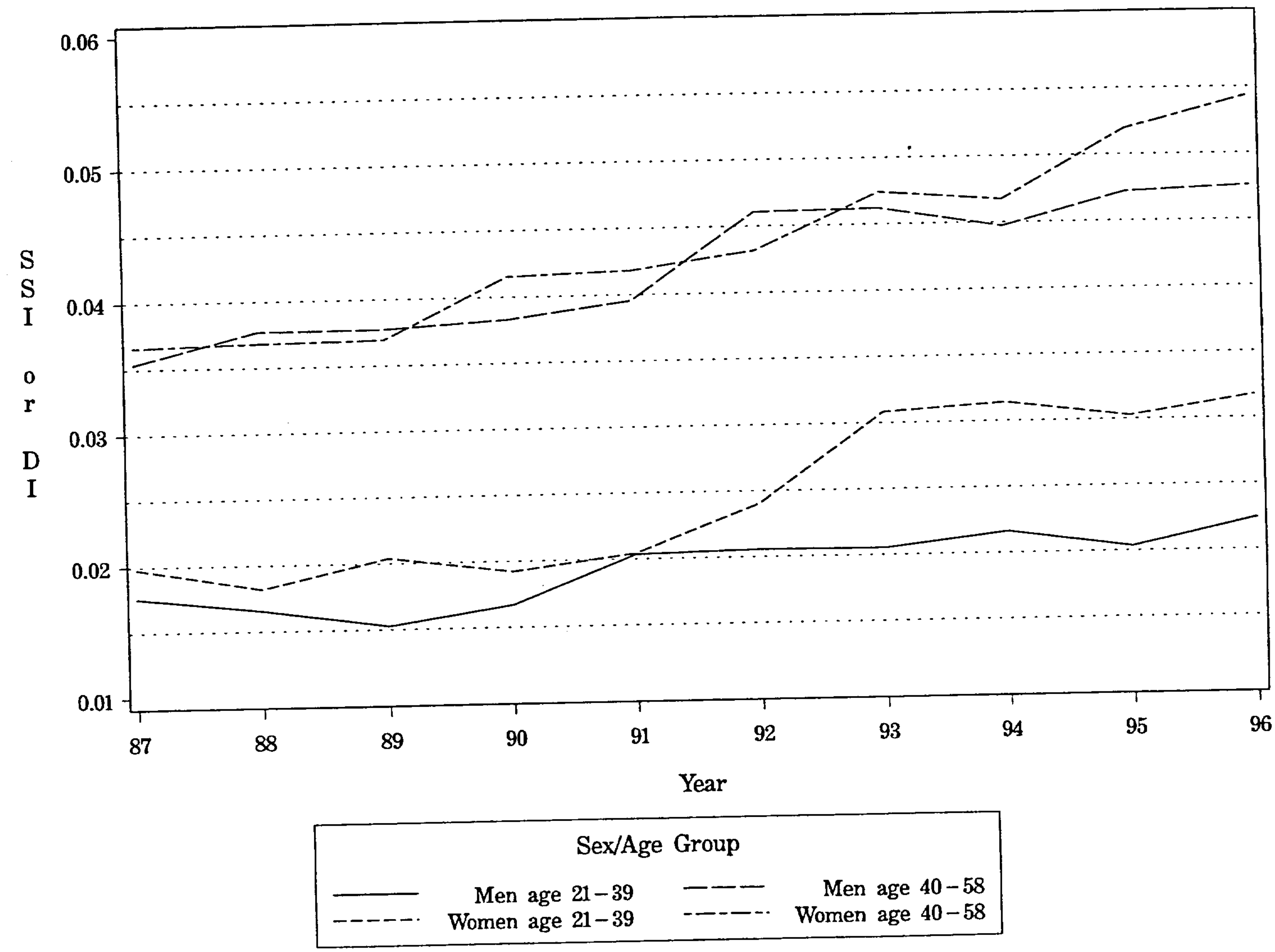

Figure B1. SSI or DI recipiency for men and women aged 21–58 in the 1988–97 March CPS.