The Impact of Work-Limiting Disability on Labor Force Participation

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Introduction

Estimates from the 2006 American Community Survey (ACS) indicate that almost eight percent of the U.S. population reports a work-limiting disability. According to the same source, the employment rate among those with work limitations is only about 20 percent. But how much of that is due to having a disability? Debate exists regarding whether respondents are accurately reporting their limitation status, which confounds the answer to this question (Bound, 1991; Bound, Schoenbaum and Waldmann, 1995; Kreider, 1999; Kapteyn, Smith and Soest, 2009). Known as the justification hypothesis, some claim that estimates of both the prevalence of disability and the unemployment rate among the disabled are upwardly biased because a subset of unemployed individuals overstate their level of disability in order to justify being unemployed. Additionally, a certain degree of underreporting of work-limitation status by the employed population may influence these statistics (Burkhauser and Houtenville, 2010).

If misreporting of disability status does indeed occur, then Census and typical survey estimates of the extent of substantially limiting disabilities and the nonemployment rate among the disabled are potentially inflated. This poses a serious problem for researchers evaluating the labor market effects of any policy aimed at the disabled or elderly populations (e.g., Supplemental Security Income, Disability Insurance, and Early Retirement). An inaccurate measure of disability in any of these cases can lead to bias not only in the impact of disability on these policies (Gordon and Blinder, 1980), but can also lead to incorrect inferences from other variables of interest, such as income (Anderson and Burkhauser, 1984; Bound, 1991; Kreider, 1999).

We contribute to the existing literature in several ways. We are the first to directly estimate the impact of disability on the labor force participation rate. For this purpose, we use data from the American Community Survey, a nationally-representative dataset that contains a rich set of economic and environmental variables, as well as several key disability measures. The ACS is advantageous to our study because we aim to evaluate reports of disability by the population of working-age individuals. The ACS is thus better suited for this research than other popular surveys[i] which have been utilized in prior studies, which limit the estimation sample to individuals over age 50. Having a base sample comprised of people whose retirement decisions are in the distant future is especially important to our analyses since disability is positively related to age, potentially affecting both the estimated impact of disability and the magnitude of reporting bias due to the justification hypothesis. Our model, which jointly estimates labor force participation and employment disability in a bivariate probit framework, produces a limitation index which can be used instead of a self-reported measure of disability.

We also estimate the impact of the environment, and its interaction with disability, on the decision to work. The rich set of variables contained within the ACS allow us to control for many important factors, such as state fixed-effects and environmental variables constructed at the local level. For this reason, we have selected this survey over the Current Population Survey and the Survey of Income and Program Participation, which do not permit the types of disability and detailed environmental comparisons in which we are interested. Using these data, our model allows us to estimate the direct impact that work-limiting disabilities have on the labor force participation rate, and how they influence the participation rate by demographic groups.

Finally, we perform a series of sensitivity analyses on the key identifying assumption used in previous research: that employed workers accurately report their level of work limitation. We uncover evidence that this is not necessarily a valid assumption.

We find that previous studies on this topic, nearly all of which restrict the analysis sample to individuals over age 50 because of data limitations, may severely underestimate the impact of disability on labor force participation and overestimate potential reporting bias. We estimate that reporting a work disability lowers the probability of being in the labor force by at least 40 percentage points for men and at least 25 percentage points for women, four times greater than the impacts implied by our replication of previous models. Additionally, our model can generate an index of disability which can be used to reduce the potential bias of misreported disability status.

A review of the related literature is presented in Section II. Section III describes the methodology (both theoretical and empirical) employed by our study. Section IV presents the results of our model and sensitivity analyses, and Section V concludes.

Background

Few studies examine the process of disablement in conjunction with the environment, despite the fact that the two major conceptual models of disability emphasize the importance of its role in influencing the reporting of disability (Nagi, 1965; WHO, 2001). The models of Nagi (1965) and the World Health Organization's International Classification of Functioning, Disability and Health (ICF) (WHO, 2001) view disability as a dynamic process involving the interaction of a person's health condition, their personal characteristics and the physical and social environment. Changes to any of these factors can potentially impact a person's functioning and activity participation, affecting employment outcomes, as well as perceived disability status. Verbrugge and Jette (1994) outline a "disablement process" model that builds on earlier theories to emphasize social, psychological, and environmental elements that may accelerate or impede progress along the pathway to functional outcomes.

Research that controls for the environment as a mitigating or contributing factor tends to focus on a subset of the adult population, preventing a more global understanding of its impact on disability. Freedman et al. (2008) analyze the Health and Retirement Study and uncover a correlation between neighborhood conditions (the built environment) and the reduced risk of instrumental activity limitations among adults ages 55 and older. Schootman et al. (2006) use the African American Health Study and find that poor neighborhood conditions (e.g., condition of houses, amount of noise, air quality, street condition, and conditions of the yards and sidewalks in front of the participants' homes) independently contribute to the risk of developing lower-body functional limitations in middle-aged African Americans. Clarke and George (2005) reveal that older adults living in central Northern Carolina report greater independence in instrumental activities in areas with greater land-use diversity, while declining density (rural as compared with semi-urban living) is associated with self-care disability. Balfour and Kaplan (2002) find that functional loss over a one-year period among persons ages 55 and older is related to self-reported problems of the neighborhood, such as heavy traffic, limited public transportation, excessive noise, and inadequate lighting in Alameda County California.

Studies analyzing behaviors of individuals with disabilities are subject to additional criticisms regarding the types of disability indicators utilized, since the accuracy of reported variables may influence key results. Self-reported employment disability (i.e., work limitation) measures are included in many national surveys and are commonly used as a proxy for health-related or disability-related work capacity. However, these may insufficiently capture health-related work capacity, leading to an error-in-variables bias which would result in an underestimation of the impact of health-related work capacity on economic outcomes (Bound, 1991). Kirchner (1994) and Hale (2001) further assert that employed workers are less likely to report a work disability. Their findings suggest that work disability measures may be sensitive to the economy—states with reduced employment would have higher disability prevalence rates. This is of less concern for surveys that have multiple disability questions capable of capturing respondents with disabilities who deny having any work limitations.

The appropriateness of using work disability as a proxy for health/disability-related work capacity is also suspect because some people who are not working may falsely report a health-related work disability to justify not working (to themselves or others, consciously or unconsciously). This is known as the justification hypothesis. If the justification hypothesis is true, then any measures that incorporate a work disability question will *overstate* the population with disabilities and *understate* the employment of people with disabilities.

Borrowing from the social psychology literature, the justification hypothesis can be viewed as a form of cognitive dissonance, which was first introduced by Festinger (1957). Particularly in America, working hard is seen as a badge of honor, while the consistently unemployed are often viewed as being lazy and a drain on society. Given these social norms, the first hypothesis of the theory of cognitive dissonance suggests that people who are unemployed will attempt to rationalize their behavior (reduce the dissonance, or inconsistency, between their views and their observed behavior). The theory does not imply that respondents will blatantly lie, rather they will convince themselves that an injury limitation is of greater severity than in reality.

In the economics literature, a growing number of studies address the justification hypothesis by estimating the likelihood of reporting disabilities as a function of personal risk factors and household and environmental characteristics. Bound, Schoenbaum, and Waidmann (1995) find that reporting a work disability is not entirely a function of health status and functional ability but is also a function of race, educational attainment, and the nature of past work. Dwyer, Sabatini and Mitchell (1999) explore alternative measures of health, function, and disability (work disability, self-reported poor health, a health condition index, and an ADL/IADL/Functional Limitation Index). Their findings are consistent across measures, suggesting that work disability is a sound measure.

Benítez-Silva et al. (2004) use the HRS to examine the self-reported disability status of a subset of persons who had applied for disability benefits from the Social Security Administration (SSA) and determined that "overall, the individual's evaluation of their disability is on average the same as the SSA evaluation of that disability" (p. 668). Burkhauser et al. (2002) examine persons reporting a disability, but not a work limitation, in the National Health Interview Survey (NHIS). The study uncovers that of those persons who reported impairments that are consistent with limiting employment, such as being deaf in both ears or blind in both eyes, only 38 percent and 69 percent, respectively, also report being "unable to work or to be limited in the kind or amount of work they do." This suggests other factors (e.g., accommodations and the environment) may influence the reporting of disability, specifically of a work disability.

In one of the most comprehensive studies to date, Kreider (1999) develops an index of work disability by estimating work limitations as a function of an extensive list of explanatory variables using data from the first wave of the Health and Retirement Survey. By testing whether those who work and those who do not work report work limitations in the same fashion, he finds that non-workers "over-report" work limitations, which lends support to the justification hypotheses. Over-reporting is associated with being female, young, less educated, nonwhite, unmarried and a former blue-collar worker. With regard to labor market outcomes, after controlling for over-reporting of work limitations, Kreider (1999) finds that work disability still has a large, negative effect on labor force participation. Kreider and Pepper (2007) perform further analyses of the justification hypothesis, with their conclusions supporting those of Kreider (1999). The analyses we perform in this paper are most similar to the research of Kreider (1999).

Methodology

Model

We present the following system of structural equations which characterize the labor force participation and disability reporting processes:

$$W_i^* = \alpha R_i + \gamma_W D^* + \theta_W (\ln(y_{1W}) - \ln(y_{0W})) + \epsilon_i^W \tag{1}$$

$$R_i^* = \gamma_R D^* + \theta_R (\ln(y_{1R}) - \ln(y_{0R})) + \epsilon_i^R \tag{2}$$

$$D_i^* = Z_i^D \delta^D + X_i^D \beta^D + \epsilon_i^D \tag{3}$$

Equation (1) models unobserved true labor force participation, W^*, as a function of reported work limitation, *R,* latent true disability status, D^*, the difference in expected income from working and not working, and an error term. Equation (2) describes the unobserved propensity to report a work limitation as a function of true disability status, the difference in expected disability-related income from reporting and not reporting a disability, and an error term. Equation (3) models D^* as a function of observed disability characteristics (e.g., sensory, mental, and physical), *Z*, and other personal characteristics, *X*.

$$\ln(y_{1W}) = X_i^{1W} \beta^{1W} + \epsilon_i^{1W} \tag{4}$$

$$\ln(y_{0W}) = X_i^{0W} \beta^{0W} + \epsilon_i^{0W} \tag{5}$$

$$\ln(y_{1R}) = X_i^{1R} \beta^{1R} + Z_i^{1R} \delta^{1R} + \epsilon_i^{1R} \tag{6}$$

and

$$\ln(y_{0R}) = X_i^{0R}\beta^{0R} + Z_i^{0R}\delta^{0R} + \epsilon_i^{0R} = 0 \quad (7)$$

Income of those in the labor force is described by equation (4) and income of those outside the labor force (i.e., nonparticipants) is described by equation (5). In these equations, identical covariates, X^W, are incorporated that include factors affecting expected income. Finally, equations (6) and (7) describe the expected disability-related income (namely Supplemental Security Income or Disability Insurance) an individual can expect from reporting a disability. This is clearly determined in large part by observed disability characteristics, *Z* (which must be verified in order to receive any sort of government benefits), but may also be affected by other personal characteristics[ii] and state fixed-effects.[iii]

We only observe reported employment disability and labor force participation, represented by

$$R_i = \begin{cases} 0, \text{if } R_i^* \le 0 \\ 1, \text{if } R_i^* > 0 \end{cases}, \quad (8)$$

and

$$W_i = \begin{cases} 0, \text{if } W_i^* \le 0 \\ 1, \text{if } W_i^* > 0 \end{cases}. \quad (9)$$

Substituting equations (3)-(5) into (1) and (6)-(7) into (2) gives the reduced form of the model as the simultaneous estimation of the probits:

$$W_i^* = X_i^W\beta^W + \alpha D_i + Z_i^W\delta^W + \epsilon_i^W \quad (10)$$

and

$$R_i^* = X_i^R\beta^R + Z_i^R\delta^R + \epsilon_i^R \quad (11)$$

While all of the structural coefficients cannot be identified through the above model, the coefficients of greatest interest are identifiable. The structural parameter α, the effect of reported work limitation on labor force participation, is identified under a joint estimation of equations (10) and (11) to account for correlated errors. Additionally, all structural parameters in equation (3), the true disability status equation, are identified under the assumption that reported disability status is equal to true disability status. The assumption that respondents accurately report their level of work limitation is a fairly strong assumption to make in general, but seems reasonable given a sample selection procedure which will be discussed in the empirical framework below.

Data

We use the 2006 American Community Survey (ACS) Public Use Microdata Sample (PUMS), including only those respondents whose age is between 21 and 64 in the household population. This yields a total sample size of 1,668,174. Following the convention of the literature, the dependent variable in the work equation (10) is an indicator for whether an individual is in the labor force. Thus, an unemployed individual who is actively searching for a job is coded as a labor force participant. This allows us to identify the effect of disability on labor force attachment rather than just employment.[iv]

We identify individuals with work disabilities using the following question: "Because of a physical, mental, or emotional condition lasting 6 months or more, does this person have any difficulty in working at a job or business?" This measure of a work limitation is the dependent variable in our disability equation, coded as 1 for individuals who give an affirmative response and 0 for those who do not. Our model does not claim to predict or evaluate the determinants of having a disability, rather the determinants of having a disability which limits one's ability to perform activities consistent with employment.

Many people with disabilities are not limited in their ability to work. All self-reported measures have weaknesses, but the ability of individuals with disabilities to select into occupations in which their disabilities will be non-limiting, such as a desk job for those with physical restrictions, further exacerbates the potential noise in this outcome variable. In each equation we include 5 separate measures of disability as independent variables to control for the presence of sensory, physical, and mental disabilities, as well as the ability to function inside[v] and outside[vi] of the home. The inclusion of these variables attempts to account for the incidence of observable disabilities (distinct from the inherently unobservable work-limitation status), and are expected to be strong predictors of a work-limitation.

One advantage of using the ACS is the ability to control for unobserved state-level variation and local environmental conditions. We thus include a full compliment of state fixed-effects, as well as local measures of the following: the unemployment rate, the prevalence of disability, and median household income.[vii] We view this as a significant improvement over previous research, particularly because of the self-reported nature of work limitation and the undeniable effect of the environment upon whether one reports being unable to work.

Each equation also includes indicators for education level, marital status, presence of children in the household, minority status, poverty status, age, and age-squared. The only differences between the set of independent variables in equations (10) and (11) are the work limitation variable, an age spline at age 61 included in the work equation, and the prevalence of disability at the Public Use Microdata Area (PUMA) level in the disability equation. The latter two variables are designed to aid in the identification process. The age spline serves to capture the effect of early Social Security benefits on employment.

While we have information on the current or previous occupations of most individuals, we choose not to include occupational fixed effects in this model. While these variables would clearly improve the overall fit of our model, their inclusion would introduce a significant amount of endogeneity into our estimates because certain types of jobs are much less accessible to the disabled.

Empirical Framework

Jointly estimating equations (10) and (11) is equivalent to the bivariate probit model with structural shift described by Heckman (1978). Several distinct simultaneous equations models are appropriate for this particular project. Since each model makes drastically different assumptions on the underlying data-generating process, we will give a brief description of the models available to us, and both a theoretical and empirical justification for our choice.

Using the notation of equations (10) and (11), the standard bivariate probit model can be modeled as the system of equations

$$W_i^* = X_i^W \beta^W + Z_i^W \delta^W + \epsilon_i^W \qquad (12)$$

and

$$R_i^* = X_i^R \beta^R + Z_i^R \delta^R + \epsilon_i^R \qquad (13)$$

This specification assumes that the error terms are distributed bivariate normal, and thus accounts for the possibility that the two outcome variables are determined jointly. If the error terms in the two equations are uncorrelated then we could estimate them separately using standard probit techniques. This framework has the drawback that you cannot directly estimate the impact of reporting a work limitation on employment. Additionally, it is difficult to argue that an individual's work limitation status does not impact the decision to be a labor force participant.

Next, we consider a simultaneous equations framework where limitation status is on the right-hand side as a continuous unobserved latent variable (rather than as a discrete reported variable). The model becomes

$$W_i^* = X_i^W \beta^W + \alpha R_i^* + Z_i^W \delta^W + \epsilon_i^W \qquad (14)$$

and

$$R_i^* = X_i^R \beta^R + Z_i^R \delta^R + \epsilon_i^R \tag{15}$$

This is akin to the structural model in Kreider (1999). The decision to model labor force participation and employment disability using (14) and (15) rather than (10) and (11) depends on whether you believe marginal increases in the propensity to report a work disability, or the discrete shift from reporting no disability to reporting a disability, drive the participation process. Equations (14) and (15) implicitly endorse the former opinion.

This decision can be supported with suggestive empirical evidence. For instance, our data from the ACS for working-age individuals indicate that, conditional on being classified into at least one of the (relatively) objective disability categories,[viii] the male employment rate is 78 percent. Even if we condition on being in at least three of the disability categories the employment rate is still 56 percent. However, the employment rate drops to only 23 percent when we condition solely on having a work limitation. This discrete jump seems to indicate that the act of reporting a work limitation reveals far more about an individual's employment outcomes than do discrete shifts the in propensity to report a limitation (of which the disability categories are obviously the strongest predictors).

We now turn to the question of identification. Heckman (1978) argues that full rank of the regressor matrix is sufficient for identification. In this case, we are able to identify the parameters of interest from the nonlinearities in the model and the structural shift from dichotomous work limitation (another reason the large jump in employment rates mentioned above is important). On the other hand, Maddala (1983) claims that the model is not identified without an exclusion restriction (in which case we would need a variable that belongs in the limitation equation which does not belong in the participation equation). However, as has been frequently pointed out (see, for example, Wilde [2000]) this condition is only required for the case of an intercepts-only model. When even one exogenous variable is added, the model is identified.

The issue of identification is not a trivial one, and despite the arguments of Heckman (1978) and Wilde (2000), we feel that an exclusion restriction is necessary in order to interpret self-reported disability in a causal manner. We therefore include the prevalence of disabled individuals in an individual's PUMA of residence only in the work limitation equation. The intuition behind an exclusion restriction in a bivariate probit model is similar to the linear instrumental variables framework, namely that three conditions must be met: (1) it must affect the reporting of work limitation, (2) it must only affect the employment decision through its effect on work limitation, and (3) the relationship must be monotonic. In our setting, we claim that the prevalence of other disabled individuals has a direct effect on reported work limitation, but only affects labor force participation indirectly through work limitation after conditioning on all other covariates. We hypothesize that the mechanism through which the prevalence of disability affects one's decision to self-identify as disabled is through cultural/social acceptance of being disabled.

While the first and third conditions are not controversial, we must justify our exclusion restriction through an examination of several possible scenarios in which the second condition would be violated.

In their seminal work on Disability Insurance, Autor and Duggan (2003) find several explanations for the rise in the disability rolls since 1984, including reduced stringency in screening and a decline in the demand for low-skilled labor. Since we use a self-reported measure, we do not worry about differences in screening requirements. However, if the proportion of disabled individuals in a locality influences the demand for labor (such as altering the composition of jobs available in the local job market) in ways not captured by the unemployment and poverty rates, then the excludability condition would be violated.

As suggestive evidence that this is not the case, we rerun each analysis also controlling for the local proportion of various classes of workers (e.g., blue-collar and white-collar) and find no significant differences in the results. While this approach introduces other forms of endogeneity to the extent that disabled individuals locate themselves in areas with certain types of jobs which are better suited to them, we feel that it is suggestive evidence that this particular scenario is not likely biasing our results.

We would also like to point out that under our exclusion restriction, peer effects of increased disability prevalence reducing the likelihood of employment is fine, so long as it is through the mechanism of increasing the likelihood of reporting a disability. This seems likely given that we condition on both the local unemployment and poverty rates.

To address the possibility that those outside of the labor force may misreport their disability status we will also estimate a bivariate probit model with sample selection. In this framework, we assume that "true" disability status is observed only when an individual is in the labor force.

For each analysis in this paper, we will estimate the model given by equations (10) and (11), and the analogue which accounts for sample selection.[ix] While the estimates from equations without a sample selection correction may not give us accurate parameters for a true work-limitation equation, they do give us accurate parameters for a *reported* work-limitation equation. We believe these parameters still have value, in that they can provide insight into the reasons for potential reporting bias, and alert future researchers to scenarios when they should be wary of reported disability status.

Hypotheses

Using the aforementioned model, we will test a series of hypotheses. First, we will examine the impact of employment disability upon labor force participation, all else held constant, through the model outlined above. We will spend the majority of our study evaluating these results, and will compare them to findings from other models used in the literature. Also, interacting work disability with demographic characteristics, such as racial and education groups, will allow for us to test whether the impact of disability differs across these characteristics.

Additionally, we will explore whether environmental characteristics such as the unemployment rate, poverty rate, and prevalence of disability have an influence on the reporting of a work disability. In other words, we will evaluate whether the environment influences the disablement process. This will be done for both the true disability equation and the reported disability equation. We hypothesize that environmental conditions will not have any effect on the true disability measure, but will likely have a significant impact on reported disability for the unemployed.

Results

Summary Statistics

Table 1 presents summary statistics for our 2006 ACS data. When compared against summary statistics from studies which use the HRS, such as Kreider (1999), these numbers illustrate one of the most impressive contributions of our study: generalizability of the data. For instance, work limitation is reported by about 7.5 percent of men in our sample. In the 1992 HRS sample used in Kreider (1999), which is restricted to those over age 50, work limitation was reported by about 22 percent of men. When examining only those outside the labor force, reported work limitation is 38 percent in our sample and 61 percent in Kreider (1999). Discrepancies persist between the data sources when we limit the ACS data to only those aged 50 and over. For instance, only 13 percent of males over age 50 in the ACS report a work limiting disability. We attribute these differences to the small (relatively) sample size of the HRS, compositional changes in the work/disability characteristics of people in older age groups from 1992 to 2006, or differences due to survey methodology[x]. Compositional changes may occur as the result of medical innovations, legislation,[xi] and changes in social norms. For instance, given the substantial increase in labor force participation by women in the 1970s, women over 50 at the time of the HRS sample may have a much different structure of labor force attachment than women over 50 today. These differences could bias coefficients and predicted values in any number of ways. Hence, we believe our sample offers a substantial improvement over those used in the existing literature.

Turning to the disability measures, we see that physical disabilities are the most commonly-reported limitation, cited by 7.7 and 8.6 percent of men and women, respectively. About 4.5 percent of both genders reported having a mental disability, while 3.5 and 2.4 percent of men and women, respectively, reported having a sensory disability.

Results

Table 2 presents the results from a bivariate probit estimation of equations (10) and (11) taking into account sample selection. We call this the "true" model since the disability equation is estimated only for those who are in the labor force. Separate analyses are run for men and women to account for systematic differences in labor market experiences across gender.

While the coefficient estimates are not very informative in any probit equation, they are of even less use in a bivariate probit where the conditional densities of the two equations are not identical (different independent variables). In this framework, a variable can have a positive coefficient in the labor force participation equation, and a negative marginal effect on participation. This would happen if that variable also has a large enough positive coefficient in the work limitation equation, in which case the variable would have a positive direct effect and a negative indirect effect (because of the negative coefficient on work limitation). Thus these coefficients should only be examined for statistical significance.

Table 3 presents more interpretable marginal effects, which take into account both direct and indirect effects on each dependent variable.[xii] The interpretation of the marginal effect of each discrete variable in Table 3 (all except for disability prevalence, unemployment, age, and poverty rate) is that if the reported value changed from 0 to 1 for every individual, then the effect on the economy would be the coefficient in Table 3. For example, if the economy went from having no individuals with a work limitation to having a completely work-limited economy, then the male labor force participation rate would drop by 39.8 percent and the female labor force participation rate would drop by 25.1 percent. Taking into account the reported levels of disability in the economy (7.5 percent for men and 7.9 percent for women), these results suggest that work-limiting disabilities decrease the male and female labor force participation rates by 3.0 percentage points and 2.0 percentage points, respectively.[xiii] This means that the observed labor force participation rates of working-aged men and women in the economy of approximately 84 and 71 percent would fall to 81 and 69 percent, respectively, under these conditions.

Examining the impact of the five disability variables included in our analysis, we find that all are strong positive predictors of work limitation. In terms of the participation equation, the three objective disability categories are strong negative predictors of being in the labor force. The other two, go-outside-home and self-care, however, have a marginal effect of close to zero. Since these variables primarily measure the severity of a disability already reported (nearly every person who gave a positive response to the go-outside-home and self-care questions also provided a positive response to at least one of the sensory, mental, or physical questions), this is suggestive evidence that the driving force behind lower labor force participation rates for the disabled is the mere existence of a disability rather than the degree of disablement. It should be noted that because of our large sample size, standard errors on these marginal effects are naturally going to be quite small. Therefore, we must evaluate each variable's economic impact rather than its statistical significance.

For both men and women, we find that among the three main disability categories, physical impairments lead to the steepest drop in employment and the greatest likelihood of having a work limitation. Mental disorders are next in both categories, while sensory impairments seem to have very little impact on employment or work limitation status.

Table 4 presents the marginal effects for the "biased" model, which assumes that everyone in our sample is accurately reporting their limitation status. A quick comparison of the effects between Tables 3 and 4 reveals the sharp differences in the impact of work-limiting disability and in the factors which influence it. For instance, the impact of disability on the labor force participation rate jumped to 66.4 percent and 56.3 percent in a completely disabled economy (i.e., one in which everyone reported having a work limitation). This translates to a drop of 5.0 and 4.4 percentage points for the male and female participation rates in our economy, respectively. When compared with the previously mentioned 3.0 percent and 2.0 percent, these inflated numbers illustrate the importance of our sample selection correction.

The environmental variable which appears to have the greatest impact on the reporting of disability is the prevalence of other disabled individuals in the respondent's PUMA of residence. This is the only environmental variable which is both statistically and economically significant in our specifications. Conforming to our hypothesis, the prevalence of disability is positively related to the reporting of disability in our "biased" model, but has no effect in the "true" model. We see this as both evidence for the justification hypothesis and indicative that our sample selection correction is successfully purging our estimates of a significant portion of this bias.

Table 5 reveals the differential impact of disability on various demographic groups. The estimates are relative to the base estimate (for men, it is -39.8) of the impact of disability on employment. Only small differences exist across male demographic groups, with Hispanic men being slightly less likely to be employed conditional on reporting a work limitation, and black men slightly more likely to be employed conditional on reporting a work limitation. The effects are much more heterogeneous and pronounced for women, however. In particular, for women with higher levels of education, reporting a work limitation appears to reduce the probability of employment by 6 to 9 percentage points.

Finally, we also run the same analysis as Kreider (1999) on our ACS data, limiting the sample to those over age 50 to be consistent with the Kreider sample. As shown in Table 6, sharp differences exist between our results and those obtained using Kreider's methods. These differences are due in part to the model (i.e., no state fixed effects and only a regional measure of unemployment), but are mostly due to the drastically different disability profile of those over age 50 as compared to the entire population.

Our results indicate that estimates of the impact of disability on work using Kreider's model and a sample restricted to the elderly may significantly underestimate its effect, as well as overestimate the reporting bias. For men, the impact of disability under these sample and model restrictions was less than 20 percent of that using our full sample and more detailed model. The estimated bias (the difference between the marginal effect of limitation in the true versus the biased models) is 50 percent greater under the restricted analysis. For women, the impact of disability under the restricted model is one third of that using the full model, and the bias is 25 percent larger.

Taking into account the observed levels of reported disability, our implementation of Kreider's age-restricted model implies that work-limiting disabilities account for only a 0.6 percent reduction in the labor force participation rate for both men and women. Generalizing this for the working-age population would mean that, according to the latest Census estimates,[xiv] at least 1.16 million individuals are not employed or searching for a job because of their disability. This is in sharp contrast to the 4.83 million people out of the labor force due to a disability implied by estimates from our more detailed model that utilizes data covering a comprehensive sample of working-age Americans (ages 21 to 64). We feel that these results illustrate the strongest justification for the importance of our study.

Sensitivity Analysis

As we mentioned before, there is also the possibility that our results are being influenced in part by those in the labor force underreporting their true limitation status. To the best of our knowledge, no prior economic studies have attempted to address this concern. Taking this possibility into account, our estimates of the impact of work-limitation on the economy can then be interpreted as a lower bound for that effect.

While we have no way to directly test the assumption that labor force participants report their limitation status accurately, we can make a range of assumptions on the type of workers who are likely to have underreported their level of disability. Table 7 presents estimates of the impact of disability on labor force participation under various assumptions about the degree of underreporting by the employed, and how to correct for it.

Assuming that all respondents who report a self-care disability (a disability which limits your ability to perform tasks around the home such as bathing) also have a work limitation implies that reporting a disability reduces the likelihood of participation in the labor force by 50 to 53 percentage points for men and 40 to 42 percentage points for women. These numbers represent an approximate upper bound for the impact of disability on employment.

The assumption that all individuals with a self-care disability is likely too strong,[xv] and thus we turn to a parametric approach to evaluate the robustness of our results. We obtain the propensity score associated with reporting a work limitation from an estimation of equation (11). We then compare the impact of disability under varying assumptions regarding the propensity score.

Intuitively, the most sound assumption of those listed in Table 7 is that individuals with a limitation propensity score greater than 0.9 be treated as having a work limitation. In this case, the impact of disability on employment is a reduction of 48 to 49 percentage points for men and a decrease of 35 percentage points for women.

Ultimately, if underreporting of work-limitation status occurs among labor force participants, it is impossible to isolate the true effect of disability on the economy without knowing specifically who is reporting inaccurately. The reader is left to draw their own conclusions based on which set of assumptions they believe.

Conclusion

Using a bivariate probit framework, we attempt to estimate the impact of work-limiting disability on labor force participation. Accounting for the justification hypothesis using a sample selection correction, we find that work-limiting disabilities reduce the percentage of men and women in the labor force by at least 3.0 percent and 2.0

percentage points respectively, or about 4.8 million total individuals in the American economy. These estimates represent a lower bound if it is assumed that there is some degree of underreporting of disability among labor force participants.

We have improved on previous research primarily through the use of newly available nationally-representative data, which allows us to control for unobservable state-level variables and local environmental variables. Our model implies that disability is a far more serious problem than past research on this topic would seem to indicate. Our replications suggest previous models would conclude disability is responsible for only a 0.6 percentage point decline in labor force participation for each gender, or about 1.2 million total individuals. Our model implies the impact of disability is at least 4 times greater. The sharp contrast between these results demonstrates that the disability index generated by our model is likely a much better predictor of true disability status than previous indices

The differences between our results and those of prior studies stems primarily from differences in the sample composition (we use a nationally-representative survey covering individuals ages 21 to 64, as opposed to restricting to older Americans close to retirement age) and the precision of certain control and environmental variables (i.e., including state fixed-effects versus regional fixed-effects, and environmental variables at the local level). Additionally, taking into account possible underreporting by the employed population produces an even larger gap between our estimates and those of previous studies.

We find moderate differences in the impact of disability across various demographic characteristics. Additionally, we find that environmental variables such the prevalence of disability in one's locality have a significant effect on the reporting of work limitations.

The large disparities between the bivariate probit results with and without a sample selection correction imply that the justification hypothesis may account for a significant portion of bias in survey responses. This is still true even under the most liberal assumptions of underreporting among those in the labor force. This underscores the importance of the limitation index which can be derived from our model, which should correct for bias in self-reports of employment disability in empirical work. In future work, we hope to focus more closely on the justification hypothesis, estimating the magnitude and determinants of the reporting bias.

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Table 1:
Descriptive Statistics

	Male		Female	
	Mean	Std Deviation	Mean	Std Deviation
In Labor Force	0.846	0.361	0.711	0.453
Work Limitation	0.075	0.263	0.079	0.270
Employed	0.813	0.390	0.679	0.467
Blue Collar	0.259	0.438	0.047	0.212
Age	43.129	12.024	43.242	12.011
Black	0.082	0.275	0.102	0.302
Hispanic	0.117	0.322	0.111	0.315
Other	0.063	0.242	0.068	0.251
No HS Degree	0.124	0.330	0.105	0.306
High School Graduate	0.293	0.455	0.276	0.447
Some College	0.287	0.452	0.319	0.466
College Graduate	0.188	0.391	0.197	0.398
Post Graduate	0.108	0.310	0.104	0.305
Presence of Child Under 18	0.415	0.493	0.454	0.498
Married	0.639	0.480	0.624	0.484
Sensory Disability	0.035	0.184	0.024	0.153
Mental Disability	0.044	0.206	0.045	0.208
Physical Disability	0.077	0.267	0.086	0.281
Self-Care Disability	0.021	0.142	0.025	0.155
Go-Outside-Home Disability	0.028	0.166	0.036	0.187
Observations	802,023		866,151	

Data are from the household population of the 2006 American Community Survey, and are restricted to respondents ages 21 to 64.

Table 2:
Bivariate Probit with Sample Selection Results (True Model)

	Men		Women	
	Limitation	Employment	Limitation	Employment
Employment Disability	-	-2.216***	-	-2.040***
	-	(0.014)	-	(0.012)
Sensory Disability	0.093***	0.009	0.063*	-0.011
	(0.029)	(0.014)	(0.032)	(0.015)
Mental Disability	0.841***	-0.051***	0.750***	-0.000
	(0.021)	(0.014)	(0.020)	(0.013)
Physical Disability	1.151***	-0.165***	1.002***	-0.011
	(0.015)	(0.012)	(0.014)	(0.009)
Self-Care Disability	0.106**	-0.024	0.123***	-0.018
	(0.045)	(0.024)	(0.037)	(0.020)
Go-Outside-Home Disability	1.157***	0.166***	1.052***	0.116***
	(0.035)	(0.021)	(0.027)	(0.018)
Age	0.057***	0.127***	0.068***	0.108***
	(0.004)	(0.002)	(0.004)	(0.001)
Age Squared	-0.001***	-0.002***	-0.001***	-0.002***
	(0.000)	(0.000)	(0.000)	(0.000)
Disability Prevalence	2.184***	-	1.713***	-
	(0.326)	-	(0.350)	-
Age Spline at 61	-	-0.143***	-	-0.150***
	-	(0.005)	-	(0.005)
Black	-0.032	-0.165***	0.051***	0.121***
	(0.022)	(0.009)	(0.019)	(0.007)
Hispanic	-0.071***	0.0827***	-0.068***	-0.049***
	(0.019)	(0.009)	(0.019)	(0.007)
Other	0.008	-0.240***	-0.006	-0.151***
	(0.022)	(0.010)	(0.021)	(0.007)
High School	-0.042**	0.117***	0.113***	0.283***
	(0.018)	(0.008)	(0.021)	(0.007)
Some College	-0.044**	0.153***	0.132***	0.407***
	(0.018)	(0.008)	(0.020)	(0.007)
College Degree	-0.185***	0.265***	0.010	0.396***
	(0.021)	(0.009)	(0.022)	(0.008)
Postgraduate	-0.264***	0.315***	0.021	0.554***
	(0.024)	(0.010)	(0.025)	(0.008)
Presence of Child Under 18	-0.052***	0.094***	-0.118***	-0.259***
	(0.013)	(0.006)	(0.012)	(0.005)
Married	-0.094***	0.180***	-0.235***	-0.575***
	(0.012)	(0.006)	(0.012)	(0.005)
Unemployment Rate	0.732	-0.120	-0.019	-0.859***

Table 2:
Bivariate Probit with Sample Selection Results (True Model)

	(0.702)	(0.264)	(0.562)	(0.198)
Poverty Rate	-0.751***	-0.563***	-0.355	0.119*
	(0.213)	(0.080)	(0.253)	(0.063)
Poverty Status	-0.146***	-0.862***	-0.186***	-0.822***
	(0.024)	(0.008)	(0.020)	(0.007)
Constant	-3.703***	-0.773***	-3.750***	-0.728***
	(0.122)	(0.066)	(0.160)	(0.055)
Rho	0.98		0.98	
Observations	802,023	802,023	866,151	866,151

Robust Standard Errors in parenthesis. ***, **, and * indicate significance at the 0.01, 0.05, and 0.10 levels, respectively.
The True model refers to the bivariate probit with sample selection specification, which assumes that only individuals in the labor force accurately report work limitation.

Table 3:
Marginal Effects from "True" Model

	Men		Women	
	Employment	Limitation	Employment	Limitation
Employment Disability	-0.398		-0.251	
Disability Prevalence	-0.000	0.001	-0.000	0.000
Unemployment Rate	-0.000	0.000	-0.001	-0.000
Poverty Rate	-0.001	-0.000	0.000	-0.000
Age	-0.006	-0.001	-0.004	-0.000
Black	-0.029	-0.001	0.033	0.002
Hispanic	0.014	-0.002	-0.0134	-0.002
Other	-0.045	0.000	-0.045	-0.000
High School	0.020	-0.002	0.076	0.004
Some College	0.025	-0.002	0.110	0.005
College Degree	0.043	-0.006	0.106	0.000
Postgraduate	0.050	-0.008	0.138	0.001
Presence of Child Under 18	0.016	-0.002	-0.073	-0.004
Married	0.032	-0.003	-0.153	-0.008
Mental Disability	-0.025	0.055	-0.015	0.043
Physical Disability	-0.049	0.089	-0.018	0.0629
Sensory Disability	0.000	0.004	-0.004	0.002
Self-Care Disability	-0.006	0.004	-0.007	0.005
Go-Outside-Home Disability	-0.001	0.102	0.008	0.078
Poverty Status	-0.203	-0.005	-0.274	-0.006

The True model refers to the bivariate probit with sample selection specification, which assumes that only individuals in the labor force accurately report work limitation.
For all discrete variables *x*, the marginal effect is calculated by taking the difference between the conditional mean function when evaluated at x=1 and x=0.
For the unemployment rate, poverty rate, and disability prevalence, the marginal effect is calculated as the difference between the observed conditional mean function and the conditional mean function after multiplying the variables by 1.10.
For age, the marginal effect is calculated as the difference between the observed conditional mean function and the conditional mean function after adding 1.

Table 4:
Marginal Effects from "Biased" Model

	Men		Women	
	Employment	Limitation	Employment	Limitation
Employment Disability	-0.664	-	-0.563	-
Disability Prevalence	-0.001	0.001	-0.001	0.002
Unemployment Rate	-0.000	0.000	-0.001	-0.000
Poverty Rate	-0.001	-0.001	0.001	-0.001
Age	-0.006	0.001	-0.004	0.000
Black	-0.032	0.005	0.032	0.001
Hispanic	0.018	-0.009	-0.009	-0.009
Other	-0.045	-0.000	-0.041	-0.004
High School	0.0233	-0.009	0.079	-0.007
Some College	0.0322	-0.014	0.115	-0.013
College Degree	0.051	-0.023	0.114	-0.022
Postgraduate	0.059	-0.027	0.146	-0.024
Presence of Child Under 18	0.017	-0.004	-0.070	-0.005
Married	0.037	-0.014	-0.147	-0.008
Mental Disability	-0.121	0.169	-0.108	0.158
Physical Disability	-0.203	0.282	-0.142	0.242
Sensory Disability	-0.005	0.010	-0.013	0.012
Self-Care Disability	-0.038	0.032	-0.050	0.034
Go-Outside-Home Disability	-0.160	0.270	-0.179	0.281
Poverty Status	-0.214	0.028	-0.280	0.028

The Biased model refers to the full bivariate probit specification, which assumes that all respondents accurately report work limitation status.
For all discrete variables x, the marginal effect is calculated by taking the difference between the conditional mean function when evaluated at x=1 and x=0.
For the unemployment rate, poverty rate, and disability prevalence, the marginal effect is calculated as the difference between the observed conditional mean function and the conditional mean function after multiplying the variables by 1.10.
For age, the marginal effect is calculated as the difference between the observed conditional mean function and the conditional mean function after adding 1.

Table 5:
Differential Impact of Disability on Labor Force Participation Across Demographic Groups
("True" Model)

	Men	Women
Black	0.014	-0.035
Hispanic	-0.016	0.027
Other	0.041	0.066
High School	-0.021	-0.068
Some College	-0.005	-0.078
College Degree	-0.026	-0.069
Postgraduate	-0.027	-0.091
Presence of Child Under 18	-0.021	0.062
Married	-0.044	0.068

The numbers in this table represent the difference between the impact of disability for the indicated group and the base group (white, no high school degree, no children, and unmarried). For instance, taking the estimated impact of disability for men from Table 3 of -39.8, the first entry in the table tells us that the impact of disability for black men is less severe than for white men at only -38.4.
The True model refers to the bivariate probit with sample selection specification, which assumes that only employed individuals accurately report work limitation.

Table 6:
Impact of Disability on Labor Force Participation Under Sample and Model Restrictions

	Men		Women	
	True	Biased	True	Biased
Analysis run on comprehensive sample using our model	-0.398	-0.664	-0.251	-0.563
Analysis run on age-restricted sample using the reduced model*	-0.079	-0.465	-0.081	-0.475

*The reduced model is the closest approximation to the model used in Kreider (1999). Some variables used in the Kreider study were specific to the dataset used (HRS) and were not available in the ACS.
The Biased model refers to the full bivariate probit specification, which assumes that all respondents accurately report work limitation status.
The True model refers to the bivariate probit with sample selection specification, which assumes that only employed individuals accurately report work limitation.

Table 7:
Impact of Disability on Labor Force Participation: Accounting for Underreporting of Limitation Status by the Employed

	All individuals satisfying condition are assumed to have a work limitation		Only individuals in the labor force satisfying the condition are assumed to have a work limitation		Individuals in the labor force satisfying the condition are excluded from the analysis	
Condition	Men	Women	Men	Women	Men	Women
Respondent reports having difficulty caring for self at home	-0.507	-0.404	-0.506	-0.410	-0.532	-0.429
Work-limitation propensity score >=0.50	-0.495	-0.389	-0.500	-0.396	-0.549	-0.439
Work-limitation propensity score >=0.75	-0.500	-0.382	-0.505	-0.387	-0.531	-0.413
Work-limitation propensity score >=0.90	-0.488	-0.354	-0.490	-0.352	-0.500	-0.359

*Three separate methods of adjusting the model were attempted to account for possible underreporting by the employed population. First, we changed all observations of the limitation variable to 1 which satisfied the given condition. Second, we changed only those observations corresponding to the employed population which satisfied the given condition. Finally, we dropped all employed individuals satisfying the given condition. The propensity score was derived from estimating the same limitation equation presented in the text.

[i] The Health and Retirement Study (HRS) is often used by those performing disability-related research covering older Americans. Using the ACS instead of the HRS gives us robust demographic data covering the comprehensive set of working-age individuals at the cost of detailed disability-specific information. However, given the focus of this paper, and the complex data-hungry techniques used, the ACS is preferable.

[ii]Government assistance programs such as Supplemental Security Income (SSI) only provide support if an individual falls below certain asset thresholds, so personal characteristics can proxy for financial eligibility determinants.

[iii]The application process is at the state level, with certain states known to be more or less stringent in granting assistance.

[iv] All analyses were rerun using employment instead of the labor force indicator, with no substantive differences observed. As with all analyses not presented in this paper, results are available upon request.

[v] Includes the inability to dress, bathe, or get around inside the home.

[vi] Includes the inability to leave the home alone for basic functions such as shopping or visiting the doctor.

[vii] The local variables were constructed at the PUMA level, a geographical division of the country into approximately 630 localities.

[viii] A respondent is classified by an affirmative response to questions regarding having a sensory, physical, or mental disability. We do not include the self-care or go-outside-home disabilities in constructing the cited statistics.

[ix] All analyses are run in both Stata and LIMDEP as a robustness check to ensure our results are not sensitive to the convergence algorithm.

[x] Small differences in the wording of questions, or even differences in the aim of the survey itself (The HRS is clearly focused on health while the ACS is not) can lead to different responses.

[xi] Legislation such as the Americans with Disabilities Act (ADA) and the Age Discrimination in Employment Act (ADEA).

[xii] These effects were computed using the unconditional mean function for bivariate normal dependent variables. For example, to compute the marginal effect on employment, the unconditional mean function would be $\Phi_2(X^W\beta^W+\alpha, X^D\beta^D, \rho)+\Phi_2(X^W\beta^W, -X^D\beta^D, -\rho)$. The figures presented are the average of each individual's marginal effect.

[xiii] These calculations are derived for men and women as follows: 0.075*0.398 = 0.03; 0.079*0.251 = 0.02.

[xiv] This calculation is based on the 2009 Census estimate of the working-age population of 193 million people.

[xv] For instance, this could include an individual confined to a wheelchair, who is still capable of holding many jobs.