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By John Bound

To study behavioral responses to social insurance programs, researchers have often used replacement rates as explanatory variables in regression analysis. Yet since the replacement ratio is endogenous, doing so is problematic. Researchers analyzing unemployment insurance or workers' compensation have surmounted this issue by capitalizing on the fact that variation in benefits across states or across time within states generates variation in the replacement ratio. This plausibly exogeneous variation can be used to identify causal effects. Since disability insurance (DI) is a nationally run program, exogenous cross-state variation in benefits does not exist. This fact does not justify ignoring the endogeneity of the replacement ratio. Rather, it suggests that one must look to other sources of information if one is to draw inferences about the effect of DI on work-force attachment.

In "The Health and Earnings of Rejected Disability Insurance Applicants" (Bound, 1989), I drew inferences about the labor-market potential of DI beneficiaries from information on the labor-market activity of DI applicants who failed to pass the medical screening necessary to qualify for benefits. Presuming that rejected applicants are in better health than beneficiaries, I assumed that their labor-market performance provided an upper bound for that which could be expected of beneficiaries, were they not on DI. In his comment, Donald O. Parsons (1991) disputes this claim. He points to a number of reasons for thinking that the behavior of rejected disability-insurance applicants is not comparable to what it would have been had they not applied for DI benefits. He argues, therefore, that it contains no information on the disincentive effects of DI.

Parsons lists three ways that DI might influence the labor-market activity of denied applicants: denied applicants may be out of the labor force while awaiting appeal; they may be out while planning to reapply; or they may face increased obstacles in returning to work, owing to the period of time they were out of work while applying for benefits. I acknowledge each of these possibilities in my original article but after some argument claim, "Although each of the effects mentioned is probably real, it is hard to imagine that they, rather than ill health, provide the major explanation for the low earnings and labor force attachment of rejected disability insurance applicants" (p. 492). The disagreement between Parsons and me does not concern whether the effects Parsons mentions are real, but how important they are for explaining the weak labor-force attachment of denied DI claimants.

Here, I address each of the three issues raised by Parsons, expanding on what I wrote in section II-A of "Health and Earnings." Before doing so, it is important to explain why it was that I used the information on rejected applicants to make inferences about the labor-supply effects of disability insurance rather than relying on more traditional regression methods such as those used by Parsons (1980a,b).

I. Why Not Use Regression Methods?

Parsons (1991) is correct in interpreting my approach to studying the disincentive effects of DI as a rejection of the more standard approach. The "standard approach" has been to employ regression techniques to compare the labor-force par-
participation rates of those with high replacement rates (those whose potential DI benefits would replace a relatively large fraction of their before-disability earnings) to the participation rates of those with low replacement rates. The difference in the participation rates between these two groups is taken to be an estimate of the impact of DI. Researchers following this strategy (Parsons, 1980a, b; Frederic B. Slade, 1984) have typically concluded that DI has had large disincentive effects, inducing a virtually one-for-one drop in participation rates.

Parsons suggests that I base my rejection of the standard approach on the logit equations I estimated on a sample of individuals who had never applied for DI benefits—estimates that show DI to have large "effects" on employment even for those men who report never having applied for benefits. This is not true. My rejection of the standard approach is based on the fact that the replacement rate is an endogenous variable. Since replacement rates for DI are decreasing functions of past earnings, it is difficult to determine whether it is generous replacement rates or low earnings that induce individuals to leave the labor force. There is ample evidence that older working-aged men who are out of the labor force increasingly tend to be less educated and in poorer health and otherwise to have lower earnings potential than their working counterparts (Parsons, 1980a, b; Chinhui Juhn, 1990; Bound and Timothy Waidmann, 1990). That the growth of the availability and generosity of disability benefits may have allowed or induced these men to leave the labor force is one possible explanation of this correlation, but not the only one. Estimates of the disincentive effects of DI that are based on using the replacement rate as an explanatory variable in participation equations are simply reproducing the negative correlation between labor-force participation and earnings. They do not help distinguish among the various explanations of this correlation.1

The logit estimates do highlight just how problematic a causal interpretation of the coefficient on the replacement ratio is. For a sample of men who identify themselves as having never applied for DI benefits, the coefficient on the replacement ratio is large and statistically significant. Parsons questions these results, pointing out that I used whether a man was employed rather than whether he was a labor-force participant as my dependent variable. In footnote 2 of his comment, Parsons suggests the possibility that the majority of nonemployed nonapplicants are actually not out of the labor force, but rather are unemployed. In fact, only 12 percent of the 14.9 percent not working report looking for work.2 To test the sensitivity of the results reported in table 6 of Bound (1989) to the use of employment rather than labor-force concepts, I reestimated the equations including those looking for work with the employed. Estimates of the coefficient on the replacement rate were uniformly larger in magnitude than the estimates reported in table 6 of Bound (1989), confirming my presumption that my original results were not driven by my specification of the dependent variable.

II. The Influence of DI on the Labor-Market Activity of Denied Applicants

A. Appeals

The calculations I reported in Bound (1989) were based on analysis of the 1972 Survey of Disabled and Non-Disabled Adults and the 1978 Survey of Disability and Work.3 I restricted my attention to men as a function of exogenous information and then incorporates these predicted values in the final estimating equations. The problem with this strategy is that it is hard to have faith in the legitimacy of the exclusion restrictions required in order to generate instruments.

1 It is not clear that these men are unemployed rather than out of the labor force, according to the official definition of this term.

2 Both surveys were done by the Bureau of the Census for the Social Security Administration. Detailed descriptions of these two surveys can be found in Users' Manual for the 1972 Survey of Disabled and Nondisabled Adults: Description and Documentation.

1 Robert Haveman and Barbara Wolfe (1984) try to avoid the endogeneity of the replacement rate by utilizing a procedure that initially predicts disability benefits.
who applied for benefits before 1971 (in the case of the 1972 survey) and prior to 1977 (in the case of the 1978 survey). Since the surveys were done during the summer months, the rejected applicants represented in my tables must have applied for benefits at least 18 months prior to the survey data. For most, the gap was substantially larger than 18 months (indeed, more than half the men represented in table 2 of Bound [1989] applied for benefits more than three years prior to the surveys). Except in the case of the very few rejected applicants who pursue their appeals into the federal district courts, the appeals process takes less than 18 months. Thus, virtually none of those represented in my tabulations on rejected applicants would have still been involved in appealing a denial.

B. Reapplication

What about men who are waiting to reapply? The social-security surveys contain no information on the number of individuals who intend to reapply for benefits. One can get some idea about how many do reapply by using tabulations reported by Ralph Treitel (1976). Studying a cohort of applicants denied benefits in 1967, Treitel (1976) reports the number receiving benefits six years later. Of those men who were initially denied benefits, 28.9 percent eventually received disability benefits some time during the six years that followed their applications. This may seem like a large number. However, it is important to note that the majority of those initially rejected who eventually qualify for DI do so through the appeals process and not by reapplying. As I argued above, virtually all of those who successfully appeal have been eliminated from the data I analyzed. Among those individuals who did not appeal, 15.5 percent eventually received disability benefits. Presumably, a higher fraction must have reapplied for DI benefits. However, the implications of these statistics for the proportion of a cross section of rejected applicants who might be planning to reapply for benefits is unclear. In any cross section, such as my sample, some of those who would eventually reapply for benefits have done so already. Either they have succeeded in their reapplication effort, in which case they would no longer be in the rejected pool, or they have failed and have perhaps given up hope of ever succeeding.

Assuming that half of those who reapply pass the screening on the second round but that half of those who will reapply have already done so, I conclude that approximately 15 percent of the rejected applicants represented in table 2 of my 1989 paper may still be considering reapplication. To the extent that these men are less likely to work than they would have been were they not hoping to qualify for DI benefits, their inclusion will bias estimates of the labor-force attachment of the rejected applicants. However, there is reason to believe that this bias is not large. Those men who are likely to take the gamble that they might qualify for DI with a second application are, presumably, those for whom the opportunity costs of doing so are low: those out of work anyway and those who perceive their chance of succeeding to be high (i.e., those in the worst health).

Suppose that 15 percent of those in my two samples are planning to reapply for benefits. Suppose further that, were it not for this fact, they would be half as likely as other rejected applicants to work, but as it is, they do not work. A little algebra shows that the presence of these hopefuls biases downward the employment rates among the rejected pool by 8 percent. Indeed, even under the extreme assumptions that the 15 percent who reapply would have been no
less likely to return to work in the absence of their reapplication than those who do not reapply and that, while waiting to reapply, they do not work at all, the downward bias on the employment rate is only 15 percent.

I conclude from these calculations that, although some rejected applicants may purposely remain out of work in order to strengthen their case for reapplication, this is the case with very few among the pool of rejected applicants. Their presence in the sample should not have a large effect on the employment rate for the entire group.

C. Treitel's Work

Up to this point, I have argued that the effect of possible appeals and reapplications on the observed employment rates of the pool of rejected applicants represented in table 2 of Bound (1989) is small. An alternative strategy, which I followed in my original (1989) “Health and Earnings” paper and which Parsons (1991) followed in his comment, uses information Treitel (1976) reports to construct “revised” estimates of the labor-force attachment of rejected applicants—estimates that net out those who successfully appealed or reapplied.

One problem with using Treitel's (1976) tabulations for the purpose of making inferences about the labor-force potential of DI beneficiaries should be noted from the start. To draw inferences about the labor-market potential of the stock of DI beneficiaries, the natural comparison group is the stock of rejected applicants: all rejected applicants between the ages of 45 and 64 who are currently living, not a single cohort five or six years after their original applications. Unlike a single cohort several years after rejection, a cross section of rejected applicants contains some who applied in the very recent past and others who applied much longer ago. The trouble with trying to infer what is true about a cross section from data on a cohort can be clarified with an extreme example. Suppose that two kinds of men are denied disability benefits: those in poor health who are incapable of work and suffer high mortality and those who are able-bodied and enjoy low mortality rates. Even if the population of denied applicants is, in fact, dominated by the truly disabled, as long as a cohort is followed for a sufficiently long period of time, one would find that most of those left would be able-bodied. Focusing on a group of applicants five or six years after they were denied benefits might very well show a large portion of the surviving population working, but this would not be reflective of a typical cross section of rejected applicants, which would include more recent rejects. For these reasons, I did not think it appropriate to use Treitel's (1976) numbers to estimate the proportion of the rejected applicants working. Rather, I used Treitel's (1976) work heuristically and emphasized calculations based on the 1972 and 1978 surveys.

When reporting calculations based on Treitel's (1976) tabulations, I restricted my sample of rejected applicants to those who had neither appealed their decisions nor successfully reapplied. In addition to appellants and successful reapplicants, Parsons (1991) excludes those men who receive social-security retirement benefits or are dead. In so doing, Parsons has eliminated those who are the oldest and the least capable of work. Treitel (1976) separately reports results for men under the age of 50 as of 1967, men aged 50–59, and men aged 60–64. Parsons performs his calculations first on the full sample and then on the sample of men under the age of 50. The first calculation is based on a sample that contains many men who, five years after applying for DI benefits, are over the age of 65, while the sample of those under the age of 50 is skewed toward younger ages. A fairer calculation to approximate my 45–64-year-old samples as closely as possible using Treitel's (1976) data would be to use men who were under the age of 60 as of 1967. Five years later, these men will be under the age of 65. Of those men in Treitel's (1976) sample under the age of 60 in 1967, 41.6 percent of those originally denied benefits show some positive social-security earnings five years later, another 11 percent are dead, 5.2 percent are receiving DI, and 22 percent are alive but not working in covered employment. If one excludes those receiving DI benefits, 52 percent of the remainder are
working. Excluding those who died as well, 60 percent of those who remain are working. Thus, even including men under the age of 45 and excluding those who have died (both of which are exclusions that will skew the sample toward those more capable of work), the population working does not rise substantially above the 45-percent employment rate reported in table 2 of “Health and Earnings.”

Parsons would have me continue these calculations by first excluding those who are receiving social-security retirement benefits and then increasing this number by 13 percent to take account of those working for the government and, therefore, not in social-security-covered employment. Eliminating those who receive retirement benefits seems odd. If one is interested in isolating the effect of DI on labor-force participation, it is important to take into account whatever other options are available to these men. It also seems implausible to me that men who have spent most of their adult lives working in the private sector would shift to the public sector after becoming disabled.

Another feature of Treitel's (1976) numbers is worthy of note. Treitel reports whether a man worked at all in social-security-covered employment five years after having been denied DI benefits. While, as Parsons notes, these numbers will underestimate the fraction who work at all since not all employment is covered, he ignores the fact that they will overestimate the fraction working at a point in time, since part-year workers are included as having worked. This latter effect is far from trivial. Computations based on the 1972 and 1978 surveys of the disabled show the fraction of rejected applicants who worked some time during the year to be 40-percent higher than the fraction working at the time of the surveys (see table 2 in my original [1989] paper).

D. Processing Delays

Parsons implies that it is important for my argument “that rejected claimants (who are older and not in good health) face no special problems in returning to the work force after extended absence” (p. 1424). I neither believe this statement nor see it as germane to my argument. To the extent that rejected applicants face troubles returning to the work force due to their age, their health, or the time they spent out of work, so would beneficiaries. The real question is: to what extent are the delays caused by applying for disability benefits responsible for the rejected applicants' trouble returning to the work force? To evaluate this, it is important to realize that the majority of men applying for DI benefits do so after having been out of the work force for some length of time (seven months on average) and that, for most men who do apply for benefits, processing time is only a few months.

As of the early 1980's, the initial determination took an average of 46 days, reconsideration took an average of 39 days, a hearing before an administrative law judge (ALJ) took 165 days, and the appeal of the ALJ decision took 66 days. At each stage, an applicant has 60 days to appeal a decision. Using data from 1970 and 1978, I calculated the fraction of those whose claims are eventually rejected and who appeal to various levels. Results are reported in Table 1.

The numbers reported in Table 1, clarify several issues. Through the early 1970's, the vast majority of cases were decided at the initial level. Almost 80 percent of those eventually rejected would have been rejected at the initial hearing and would not have appealed. Less than 10 percent of those rejected carried their case as far as an ALJ hearing. Thus, at least as of the early 1970's, for the vast majority of individuals, processing time delayed return to the labor market by no more than 46 days. By 1975, the fraction of rejected applicants appealing had risen. Even so, almost 80 percent of those eventually denied benefits do not appeal beyond the reconsideration. Thus, it seems safe to conclude that processing delays do not contribute importantly to the time men stay out of the labor market while applying for DI benefits.

This argument rests on the presumption that the time spent out of work prior to the application would have been time spent out regardless of the existence of DI. However, one might imagine that part of the time a
TABLE 1—PERCENTAGES OF REJECTED APPLICANTS WHOSE FINAL DECISIONS ARE MADE AT EACH LEVEL

<table>
<thead>
<tr>
<th>Level</th>
<th>Percentage</th>
<th>Estimated cumulative processing delay</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>1970</td>
<td>1978</td>
</tr>
<tr>
<td>Initial hearing</td>
<td>79.0</td>
<td>70.1</td>
</tr>
<tr>
<td>Reconsideration</td>
<td>13.6</td>
<td>20.9</td>
</tr>
<tr>
<td>ALJ Hearing</td>
<td>4.2</td>
<td>4.7</td>
</tr>
<tr>
<td>Appeals Council</td>
<td>2.8</td>
<td>3.4</td>
</tr>
<tr>
<td>District courts</td>
<td>0.3</td>
<td>0.7</td>
</tr>
</tbody>
</table>

Sources: Mordechai Lando et al. (1982) and U.S. Senate Finance Committee (1982). See text for details.

A man spends out of the work force prior to applying for DI benefits could be attributed to an applicant’s interest in strengthening his case before applying. I doubt that this is the reason why applicants delay. While to qualify for DI benefits one must be out of work, there is no indication that extending the length of time one spends not working before applying for benefits lowers the probability of being denied them. In fact, Robert E. Leihy (1979) reports that those who delay filing for claims face a greater likelihood of being denied benefits.

Further indications of the potential impact of processing delays on the labor-force attachment of rejected applicants can be obtained by using Treitel’s (1976) results to compare the work-force attachment of rejected applicants who appeal the initial decision with the attachment of those who do not. Since one expects that those who appeal are in worse health than those who do not—the fact that their mortality rate is higher supports this presumption—one expects those who appeal to be less likely to work than those who do not, for this reason alone. Even so, the gap provides an upper bound on the impact of processing delays on work attachment. Focusing on the group under the age of 60, of those not receiving DI benefits, 46.2 percent of those who appealed the initial determination were working, while of those who did not, 54.6 percent were working.6 The narrowness of this gap suggests that processing delays do not have a large impact on the work-force attachment of rejected applicants.

III. Summary

As I noted in my original paper, each of the major points Parsons (1991) raises has some validity. Each suggests that applying for DI may have adversely affected the labor-market potential of even those applicants who do not pass the medical screen. Even so, it still seems likely that the labor-market performance of rejected applicants can be taken as an upper bound for the potential of the actual beneficiaries. I do not make this assertion simply on the basis of the calculations reported above. The evidence I reported in my 1989 paper from Saad Nagi’s (1969) clinical evaluation of the work capacity of applicants for DI benefits shows 33.7 percent of rejected applicants, but only 7.7 percent of beneficiaries, to be capable of work under normal conditions (see my 1989 paper for a fuller discussion of the Nagi study). The proportion of rejected applicants whom Nagi finds incapable of work is similar to the proportion of rejected applicants who, in fact, do not work. Thus, it seems reasonable to infer that the rejected applicants are in poor health and that this is the reason why many do not go back to work. Moreover, the clear contrast between the rejected applicants and the

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6Netting out those who had died gives quite similar figures: among those alive, 53.6 percent of those who appealed were working, as against 41.2 percent of those who did not appeal.
beneficiaries in their capacity to work persuasively argues that beneficiaries would be less likely to work than the rejected applicants. I was led to the conclusion that the rejected applicants’ labor-force participation was an upper bound on what could be reasonably expected of DI recipients in the absence of DI by the contrast between what I estimated to be the small effects of the DI system on the behavior of rejected applicants (see above) and the large (26 percentage points) difference in the fraction able to work according to Nagi’s analysis.

Let me emphasize, as I did in the original (1989) “Health and Earnings” paper, that the inferences I draw about the labor-market potential of DI beneficiaries based on the labor-market behavior of rejected applicants are completely consistent with the information available from other sources. In 1949, several years before DI came into existence, the Current Population Survey included a question on disability. Of men 45–54 years old, 3.7 percent were identified as disabled and unable to work, and for men 55–64 years old, this proportion was 8 percent (Marjorie E. Moore and Barkev S. Sanders, 1950). These numbers suggest that a substantial fraction of the 4.2 percent of the 45–54-year-old men and the 11.3 percent of the 55–64-year-old men on DI as of 1980 would have been out of the labor force even were they not receiving DI benefits.\(^7\)

Research that has focused on the application for DI benefits (Jonathan S. Leonard, 1979; Lando et al., 1979; Janice H. Halpern, 1979) has produced estimates that are consistent with those reported in my original “Health and Earnings” paper (see section III in Bound [1989] for a fuller discussion of these issues). Only the cross-sectional studies that focus on labor-force participation and use replacement rates as explanatory variables (Parsons, 1980a 1980b; Slade, 1984) have found large disincentive effects, and there is ample reason to doubt such estimates.

\(^7\) A fuller discussion of the historical evidence can be found in Bound (1989) and Bound and Waidmann (1990).

REFERENCES


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