# Demographic Factors in the Disability Determination Process: A Logistic Approach

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Identifying the demographic factors that affect the likelihood of being allowed disability insurance benefits under the social security program is important in assessing the operation of that program and has been the focus of continuing research. Previous studies in this area, which have been limited to cross-tabular analyses of aggregated data and the examination of a few variables at a time, have uncovered apparently sizable differences in the probability of allowance by age, sex, race, and other demographic characteristics. This article uses a much more rigorous statistical technique—a logit maximum likelihood procedure—to examine the same question. That technique permits exploration of the relationship between a particular characteristic and the probability of allowance while controlling for the effect of other characteristics considered in the analysis. The findings show that cardiovascular primary diagnoses had a higher probability of allowance than did almost all other primary diagnoses. Mobility restrictions, older age at the onset of disability, and residence in States with temporary disability programs also were associated with a higher probability of allowance. Applicants who were black or from the South were found to have a lower probability of allowance. In general, those results applied to both men and women.

In assessing the operation of the disability insurance program under the Social Security Act, it is important for analysts to identify demographic factors that affect the likelihood of benefit allowance. Though research in this area has, until now, been limited to cross-tabular analysis of aggregated data and the consideration of a few variables at a time, it has nevertheless revealed apparently sizable differences in the probability of allowance according to age, sex, race, and other demographic characteristics.

A previous article on this subject,[1] for example, which was based on data from disability applicant records for all disability decisions made in 1971, revealed that the proportion of disabled-worker applications that were allowed varied according to sex, race, and age. In particular, whites had a higher proportion of allowances than did blacks, older persons than younger ones, and men than women.

That study also attempted to determine the underlying causes of the differences in allowance rates according to sex and race. By means of cross-tabular classifications of aggregated data, it was found that differences between the applicant and insured populations in labor-force patterns, education levels, and age distributions were among the causes. More than half the differences between black and white applicants in the proportion of claims allowed could be explained by the differences in age distributions. It was also hypothesized that the lower proportion of claims allowed for black persons reflected their higher application rates since, across a distribution of severity, the probability is that the greater the number of applications the higher the proportion of less severely disabled persons.

This article examines the same subject by means of a logit maximum likelihood procedure. Instead of cross-tabular proportions, the individual applicant and the effects of independent variables on the probability of positive determination are considered. The advantages of this more rigorous technique, which permits exploration of the relationship between a particular characteristic and the probability of allowance while controlling for the affect of other variables, and the motivations for using it are numerous. Because the tables presented in the earlier article were two-way and three-way classifications, only a small number of

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[1] Mordechai E. Lando, "Demographic Characteristics of Disability Applicants: Relationship to Allowances," **Social Security Bulletin,** May 1976, pages 15-17.

variables could be considered at a time. Interpretation of the tables was therefore difficult. The omitted variables could have affected the results by obfuscating true relationships and by ignoring the variations between individuals (and the values of variables for these individuals).

Since the unit of determination is the individual, using it here as the decision unit makes sense. This article therefore *focuses on the effects of different variables on the probability* that an individual would receive a positive determination for disabled-worker benefits. By including a number of other independent variables, it was hoped that the effect of race on this process could be explored more intensively.

# Methodology

## Estimation Procedure

When probabilities are estimated by least squares, the resulting estimates, though unbiased, are inefficient. Furthermore, the random disturbances can no longer be assumed to be normally distributed, and the standard hypothesis-testing techniques are inappropriate. Because the dependent variables is dichotomous, the standard measure of predictability statistic, $R^2$, also is inappropriate.

Finally, there is no certainty that the estimated probabilities will fall within the closed [0, 1] intervals; although most probabilities should clearly do so, interpretations of those falling outside this interval are very difficult to accomplish. For these reasons, least squares is clearly an inappropriate estimation procedure.

The model presented here is estimated by a logit maximum likelihood procedure that yields consistent and efficient estimates.[2] Given $P$(probability of success) $= e^{x\beta}/1 + e^{x\beta}$ and $Q = 1/1 + e^{x\beta}$, the likelihood function [L $(\beta)$] can be written as follows:

$$L(\beta) = \overset{\pi}{t\Sigma\theta_1} \left(e^{x_t\beta}/1 + e^{x_t\beta}\right) \overset{\pi}{t\Sigma\theta_2} \left(1/1 + e^{x_t\beta}\right)$$ where

$x_t$ is a $k$-element row vector of independent variables,

$\beta$ is a $k$-element column vector of coefficients,

$\theta_1$ is the set of all observations such that a success is observed, and

$\theta_2$ is the set of all observations such that a failure is observed.

By maximizing this function, one can obtain parameter estimates. If a coefficient on a variable is positive, the interpretation is that, if that variable increases with the others held constant, then the probability of allowance increases. This specification is attractive because the logarithm of the odds ratio is a linear function of the independent variables, that is, $\ln P/Q = x\beta$. The coefficients can be interpreted as the marginal effects of a change in $x$ on this dependent variable. The negative of the expected values of the second derivatives of the logarithm of the likelihood function, evaluated at the maximum, will yield the asymptotic standard errors of the estimated coefficients. In this way, tests of significance of the independent variables can be constructed.

## The Data

*The data were taken from the 1975 Continuous Disability* History Sample (CDHS). This file contains information on applicants for disabled-worker benefits. It is a stratified random sample of all disability determinations made during the calendar year. The CDHS, however, also includes a 10-percent random subsample of all claims, which was used as the basis for this article.[3] The full sample was stratified on the acceptance-denial decision, which yielded parameter estimates that behave poorly; the 10-percent subsample was of simple random design and yielded estimates with desirable properties.

An observation had to satisfy a number of other criteria to be included in the study. Only disability claims initially determined in 1975 are considered here. This procedure was deemed most appropriate because the study was designed to focus on a particular point in time and not allow for changes in the determination process. If a person was designated to receive a partial allowance,[4] he was recorded as having received a disability determination in that year. If that decision was later amended, the date of the initial determination was dropped from the record. Because only decisions initially rendered in 1975 are relevant for the purposes of this study, 1975 amended decisions were omitted.

Claimants who met the statutory definition of blindness were also excluded. This step was taken because persons in that category have a much higher probability of receiving disabled-worker benefits. Also omitted were denials for various technical reasons not related to the merits of the claim. These reasons included: (1) Work despite impairment, (2) failure to follow treatment, (3) failure to submit to consultative examination, (4) failure to cooperate in submitting evidence, (5) failure to meet the earnings test at the alleged onset of disability or later, and (6) withdrawal of the claim. Finally, cases with impossible or missing codes for the relevant variables were also omitted.

The model presented here was run with two different data sets. In the first set, all allowances that met the above criteria were considered. In the second, only those cases in which it appeared that the claims examiner's judgment figured significantly in the determination were analyzed. Excluded on this basis were persons with allowances who were "disabled because of an impairment specifically listed in the **Listing of**

[2] See Peter Schmidt and Robert Strauss, "The Prediction of Occupation Using Multiple Logit Models," **International Economic Review,** June 1975, pages 484-485.

[3] For a detailed description of the CDHS, see Office of Research and Statistics, Social Security Administration, **Continuous Disability History Sample Restricted Use Data File: Description and Documentation—January 1978.** For the design used to decrease the sample, see appendix A to the report.

[4] When it is possible that the date of disability onset established by existing evidence may be set back by further evidence, a partial allowance is granted.

**Impairments** and medical evidence contains the specific findings listed for that impairment."[5]

The differentiation of data sets is thus based on claims examiner input. If the examiner and nonhealth factors are not considered for those who have listed impairments, the applicants concerned should not be included in the model. They are therefore excluded from the second sample. Where this decision is not so clear-cut, the applicants concerned should be included in the model. For this reason, they appear in the first sample. The number of cases excluded for various reasons and the size of the resulting data files used for analysis were as follows:

| Reason for exclusion | Number | |
|---|---|---|
| | All cases | Cases excluding those meeting medical listing exactly |
| Total in CDHS | 135,144 | 135,144 |
| In 10-percent random subsample | 106,961 | 106,961 |
| Exclusions: | | |
| Number meeting medical listing exactly | ........... | 15,028 |
| Technical denials | 6,342 | 6,342 |
| Noninitial determinations | 4,912 | 4,468 |
| Statutory blindness | 16,771 | 13,991 |
| Missing or impossible codes | 7,075 | 5,785 |
| Net cases after exclusions | 71,861 | 61,347 |

## The Model

The disability determination process is based on medical and nonmedical factors. The applicant's condition is subject to medical substantiation. The regulations, however, allow easier acceptance for older workers involved in arduous labor who are not able to find employment similar to that in which they were engaged before the onset of disability.

The model used here may be interpreted as a reduced-form equation representing the decision of the individual with a self-perceived disability to apply for benefits and the decision by the State agency with respect to entitlement to benefits. Where these effects have opposite signs, the sign of the estimated coefficient represents the dominant effect. For an example, see the following discussion of the education variable. To clarify results, two models, one for men and one for women, were estimated. The variables included in the models are age at onset of disability, education, primary diagnosis, mobility, occupation, possible eligibility under State temporary disability insurance programs, race, and residence in the South.

**Age at onset of disability.** This variable approximates the respondent's age at the onset of self-perceived disability. Although the law stipulates that successful applicants must be unable to engage in substantial gainful activity anywhere in the economy, this requirement is relaxed for older workers engaged in arduous unskilled labor to the extent that they must only be unable to return to their usual work because of significant impairments. Consequently, older workers are required to satisfy a less stringent set of requirements to receive disabled-worker benefits. This variable should therefore have a positive coefficient.

**Education.** In general, better-educated persons have more options in the job market because they possess a larger stock of general and specific human capital (earnings potential) and thus have a higher probability of engaging in substantial gainful activity. For this reason, such persons should have a lower probability of coming on the rolls. Because they do have more opportunities in the job market, however, it is expected that they would be less likely than others to apply for benefits. It is therefore anticipated that the better educated would have a tendency to be more severely disabled before making an application and thus have a higher probability of having it accepted. This effect would be especially strong in the model estimated here because no variable is included that adequately reflects the severity of the disability. The estimated sign of the education variable reflects the relative strengths of these opposite effects.

**Primary diagnosis.** The 1975 CDHS contains a four-digit code to represent the claimant's primary diagnosis—the condition primarily responsible for disablement. The standard listings were recoded into a vector of 10 dummy variables.[6] In conformance with standard procedure to satisfy constraints on the model, one dummy was omitted—in this case, cardiovascular diseases, which affected a large number of respondents. Coefficients on the other diagnoses were interpreted as deviations in the logit from that group.

**Mobility.** The model includes a three-way mobility variable created from the seven-way variable in the CDHS.[7] Lack of mobility was expected to increase the probability of disability allowance. If a degree of mobility impairment effect exists, this fact should be picked up by a significant difference in the mobility coefficients. The category "go outside without help"—which denotes little or no mobility impairment—was used as the reference group.

**Occupation.** The data set contains a three-digit occupation code that flags the claimant's predominant occupation in the preceding 10 years. By means of the standard classifications,[8] it was possible to classify occupations by one-digit codes and to include occupational dummy variables. For this classification, the code for professional, technical, and

[5] The impairment of each applicant is compared with a standard medical listing of impairments described in terms of specific symptoms, signs, and laboratory findings. If the applicant's impairment is judged severe enough to meet the listed impairment, he is presumed to be disabled and awarded benefits. Whatever examiner discretion exists in this category or in other discarded categories should be slight.

[6] For standard listings, see Department of Health, Education, and Welfare, **Eighth Revision, International Classification of Diseases (adapted)**, 1968.

[7] The CDHS has codes to designate the applicant's ability to move about: Institutionalized, confined to general hospital, bedridden (home), chairbound, housebound, go outside with help, and go outside without help. In formulating the trichotomous variable, categories 1-5 were grouped and categories 6 and 7 remained distinct.

[8] Department of Labor, **Dictionary of Occupational Titles**, vol. 2, 3d edition, 1965.

managerial occupations was arbitrarily omitted. The remaining occupational classifications are clerical and sales; service; farming, fishery, forestry, and related occupations; machine trades; bench work; structural work; and miscellaneous occupations. Some occupational groupings might have a higher probability of allowance than others. If different probabilities exist for white-collar and blue-collar workers, this factor might be reflected in the occupation coefficients because the first three groupings are more likely to be white-collar occupations and the others (with the obvious exception of the miscellaneous category) are more likely to contain a larger proportion of blue-collar workers.

**Eligibility under State temporary disability insurance.** Five States (California, Hawaii, New Jersey, New York, and Rhode Island) and Puerto Rico have insurance programs that provide workers with partial compensation for the loss of wages caused by temporary non-work-connected disability. These programs generally define disability as the inability to perform regular or customary work because of a physical or mental condition. Clearly, the eligibility requirements under these programs are less stringent than those for the social security program. Persons on the State rolls would also be more likely to apply under the social security program. Because State temporary disability insurance recipients are receiving transfer payments, they are not without income during the waiting period for disability insurance, and the psychic costs of application are minimized. Since the costs of application are minimized and the acceptance criteria are more restrictive under the social security program, such applicants should have a lower probability of acceptance. There is no reason to believe that, in general, applicants from these States would be less impaired than those from other States, however. Instead, the State disability programs may act as a screening device for the Federal program; applicants from these States can thus have a high probability of acceptance. Consequently, a dummy variable to differentiate these States was included and was expected to have a positive sign.

**Race.** The CDHS data set contained three race categories: Black, white, and other. Because of the relatively small number of applicants in the third group, it was omitted. The dummy variable was set equal to 1 if the applicant was black. In order to determine if a differential racial effect was evident in the South, a South/race interaction variable was also included.

**Residence in the South.** A dummy variable was set equal to 1 if the applicant was from the South. The variable measures whether applicants from the South have a significantly different probability of acceptance.

# Findings

## Results for Men

To see if the relationships between variables and the probability of allowance differed for men and women, the sample was stratified by sex. The results of the logit estimation of the parameters of the initial determination process for men are presented in table 1. Many of the variables exhibit a significant effect on the initial determination.

Age at onset of disability had a significant positive effect on the probability of allowance. This result is consistent with the regulations, which state that nonmedical (vocational) factors such as age, education, experience, and skills may be taken into account if a determination cannot be made on the basis of medical evidence alone.

Education did not significantly affect the probability of initial allowance. This result seems to imply that the two effects mentioned offset each other and that the net effect of increased education is zero.

Estimation of the model clearly showed that different primary diagnoses have different probabilities of initial

**Table 1.**—Logit estimation of parameter of initial determination process for all men and women, by type of variable

| Variable | Men | | Women | |
|---|---|---|---|---|
| | Coefficient | *t*-statistic | Coefficient | *t*-statistic |
| Constant | -1.3792 | [1] 6.81 | -2.0645 | [1] 7.51 |
| Age at onset of disbility | .0517 | [1] 21.29 | .0475 | [1] 14.43 |
| Education | .0099 | 1.30 | .0138 | 1.31 |
| Diagnostic group:[2] | | | | |
| Musculoskeletal | -1.1985 | [1] 17.17 | - .5020 | [1] 6.18 |
| Respiratory system | - .8053 | [1] 9.02 | - .2494 | 1.90 |
| Digestive system | -1.6775 | [1] 14.38 | -1.1340 | [1] 7.58 |
| Mental disorders | -1.0989 | [1] 12.66 | -1.5889 | 1.56 |
| Nervous system | - .2874 | [1] 2.02 | .0380 | .23 |
| Genitourinary | - .9881 | [1] 4.44 | -1.6097 | [1] 6.53 |
| Neoplasms | .4228 | [1] 3.31 | .8681 | [1] 6.59 |
| Endocrine | -1.0270 | [1] 8.56 | - .5783 | [1] 4.50 |
| Other | -1.6275 | [1] 20.55 | -1.2513 | [1] 12.11 |
| Mobility impairments: | | | | |
| Moderate | 1.2185 | [1] 10.61 | 1.7003 | [1] 8.98 |
| Severe | 1.3162 | [1] 14.63 | 1.6257 | [1] 13.86 |
| Occupation:[3] | | | | |
| None/unknown | .0398 | .31 | - .0311 | .19 |
| Clerical, sales | - .1152 | .93 | - .2112 | 1.47 |
| Service | - .1691 | 1.41 | - .2953 | [1] 2.05 |
| Farming, fishery, forestry, etc. | .1086 | .79 | - .4591 | 1.58 |
| Processing | - .6752 | .43 | - .1691 | .73 |
| Machine trades | - .1236 | .20 | - .1178 | .66 |
| Bench work | - .1005 | .74 | - .3081 | [1] 1.99 |
| Structural work | .0067 | .06 | - .2169 | .75 |
| Miscellaneous | .1356 | 1.29 | - .0595 | .35 |
| With State temporary disability insurance coverage | .1445 | [1] 2.35 | .1517 | [1] 2.15 |
| Residence and race: | | | | |
| South | .3446 | [1] 5.93 | - .1588 | [1] 2.12 |
| Race | - .6626 | [1] 7.97 | - .3216 | [1] 3.43 |
| South/race | .5069 | [1] 4.06 | .0367 | .25 |

[1] Significant at $\alpha = .05$ for a two-tail test.
[2] Excludes cardiovascular disorders here and in table 2.
[3] Excludes professional, technical, and managerial category here and in table 2.

allowance. In fact, except for neoplasms, all diagnoses had a significantly smaller effect on allowance than did cardiovascular diseases.

The model contained two mobility restriction variables to represent slight and more pronounced problems. Both of these dummy variables significantly increased the probability of initial allowance. Although the variance-covariance matrix was not available, the similarity of the coefficients seemed to indicate that they were not significantly different from each other. This result seems to imply that it is not the degree of, but the presence of, a mobility limitation that affects the probability of initial allowance.

Primary occupation was not an influencing factor in the initial determination process. Hypothesis tests showed that in none of the other occupational categories did men have a significantly different probability of allowance than did those in the reference group of professional, technical, and managerial occupations.

The effect of a State temporary disability program was to increase the probability of initial allowance. This finding seems to substantiate the hypothesis that State disability programs act as screens for the Federal program. If an applicant did not qualify for State disability insurance, he would be less likely to apply for Federal benefits.

The racial effect was measured by the use of two binary regressors. These variables, a race dummy and a South/race interaction, were employed to see if racial effects in the determination process could be limited to the South. It was found that black men had a significantly lower probability of a positive determination. Southern applicants also had a lower probability of initial allowance. The finding that the South/race interaction term was positive and significant implies that black men have a higher probability of acceptance in the South than in other parts of the country and also that the race and region effect should not be entered linearly.

The results for men not meeting the medical listings exactly were similar to those for all men, which included those who met the specifications. As table 2 shows, the only differences were a positive but insignificant coefficient on neoplasms and a positive, significant coefficient on the none/unknown occupation category. The similarity of results may tend to imply that the same characteristics hold across the two groups.

## Results for Women

The results for women in the sample (table 1) were similar to those for the men. One notable difference is that the South/race interaction was not significant. This result implies that the combination of being a black southern woman does not have an effect on the probability of allowance beyond that of being black or from the South. Of less interest are the findings that two occupational categories (service and bench work) are negative and significant and primary diagnoses of nervous conditions and mental and respiratory diseases are not significant.

As table 2 shows, the results for women not meeting the medical listings exactly differed from those for all women in the sample with respect to four primary diagnoses (mental, genitourinary, respiratory, and neoplastic diseases) and three occupations (service, bench work, and structural work). The other coefficients were similar with respect to significance. Women who did not meet the medical listings exactly differed from their male counterparts by registering an insignificant South/race interaction; they also recorded differences under two diagnoses (genitourinary and neoplasms) and two occupations (none/unknown and structural work).

## The Race Coefficient: A Discussion

The results presented here on the relationship between race and the probability of allowance are subject to some caveats. Two types of structural misspecifications could

**Table 2.**—Logit estimation of parameter of initial determination process for men and women not meeting medical listing of impairment standard, by type of variable

| Variable | Men | | Women | |
|---|---|---|---|---|
| | Coefficient | $t$-statistic | Coefficient | $t$-statistic |
| Constant | -1.9800 | [1] 10.09 | -2.3196 | [1] 8.68 |
| Age at onset of disability | .0588 | 24.55 | .0480 | [1] 14.31 |
| Education | - .0106 | 1.44 | .0129 | 1.24 |
| Primary diagnosis: | | | | |
| Musculoskeletal | -1.0161 | [1] 15.09 | - .3466 | [1] 4.39 |
| Respiratory system | - .7091 | [1] 8.02 | - .3113 | [1] 2.30 |
| Digestive system | -1.5346 | [1] 13.19 | -1.2454 | [1] 7.66 |
| Mental disorders | -1.0994 | [1] 13.02 | - .2558 | [1] 2.46 |
| Nervous system | - .6208 | [1] 4.74 | - .1506 | .90 |
| Genitourinary | -1.6415 | [1] 6.96 | -1.5516 | .42 |
| Neoplasms | .1746 | 1.35 | .1837 | 1.41 |
| Endocrine | -1.0645 | 9.09 | - .5401 | [1] 4.31 |
| Other | -1.5425 | [1] 19.87 | -1.3207 | [1] 12.46 |
| Mobility impairments: | | | | |
| Moderate | 1.0029 | [1] 11.80 | 1.3992 | [1] 12.44 |
| Severe | 1.1141 | [1] 9.83 | 1.4534 | [1] 7.88 |
| Occupation: | | | | |
| None/unknown | .2482 | [1] 2.04 | - .1985 | 1.22 |
| Clerical, sales | .0166 | .14 | - .2051 | 1.42 |
| Service | .0141 | .12 | - .1580 | 1.10 |
| Farming, fishery, forestry, etc. | .0822 | .64 | - .2845 | .97 |
| Processing | .0818 | .54 | - .1873 | .74 |
| Machine trades | .1872 | 1.75 | - .1459 | .83 |
| Bench work | - .0564 | .39 | - .2876 | 1.86 |
| Structural work | .0941 | .98 | - .6456 | [1] 2.37 |
| Miscellaneous | .0067 | .07 | - .1916 | 1.11 |
| With State temporary disability insurance coverage | .2971 | [1] 5.05 | .2698 | [1] 3.87 |
| Residence and race: | | | | |
| South | - .3078 | [1] 5.38 | - .2044 | [1] 2.78 |
| Race | - .6420 | [1] 8.10 | - .3719 | [1] 3.98 |
| South/race | .4482 | [1] 3.72 | .0194 | .13 |

[1] Significant at $\alpha = .05$ for a two-tail test.

explain the results. It may be that some of the standardizing variables do not measure what they attempt to measure. It is also possible that important underlying variables have been omitted. Both of these problems would lead to statistically inconsistent and inefficient parameter estimates.

This article may contain examples of both types. The mobility index, for example, can be considered in its own right and as a proxy for severity. In its role as an instrument for severity, it performs well if it has a high correlation with severity, in which case the estimate has desirable properties. If mobility is not acceptable as a proxy for severity, the lack of a severity measure may be a serious omission in the model. If a correlation between race and severity of impairment exists, the result would be a biased coefficient on race. Because of a lack of data, it is impossible to make statements about either the correlations between severity and mobility or severity and race.

Another variable that may involve similar problems is occupation. It may be that one-digit occupation codes are too general and therefore do not reflect changes in racial composition across occupation. If, when occupations are more finely differentiated, they have varying probabilities of allowance that are not reflected in the one-digit codes, a biased racial coefficient could result. This problem may have been minimized by means of the classifications. As noted earlier, the first three occupational groups are more likely to be composed of persons in white-collar jobs and the remainder tend to be more heavily represented by blue-collar workers. It is therefore hoped that the one-digit classifications reflect the risk differentials in different occupations. Once again, it cannot be determined whether or not this is the case.

Aside from the variables that may have been subject to measurement error, underlying regressors correlated with race may have been omitted. The earlier article hypothesized that observed racial differences in allowed claims were caused partly by the lower earnings of blacks, their higher probability of unemployment, their lower educational attainment, and their greater tendency to apply for benefits. Many of these effects were not controlled for in the cross-tabular analysis but are controlled for directly or by proxy here. Controlling for these variables implies that these differences will not be measured by the coefficient on race. Nevertheless, it may be that the earlier hypotheses are valid. A process may exist by which chronically unemployed persons would be found to be more likely to apply for disabled-worker benefits. The sample taken here, however, was created to exclude persons without sufficient quarters of coverage to be eligible for these benefits, so the chronically unemployed should not be in the sample.

Similarly, the model controls for differences in human capital, though not for differences in actual earnings, by including an education variable. Median income for blacks is known to be lower than that for whites, but it is not clear that this difference continues to exist after the controls used here are applied. If the difference remains, it would not be surprising to find, for a given level of severity, more black applicants because of their higher earnings replacement rate. This result would imply that blacks have a higher proportion of marginally disabled applicants and a lower probability of acceptance.

If the human capital variable does not act as a good instrument for income, the lower median income of blacks could operate through a number of mechanisms to induce marginally disabled persons to apply for benefits. Through networks such as national welfare rights organizations, poor persons are more likely than others to be privy to information about the program and thus are more likely to apply for benefits. Furthermore, welfare recipients may perceive a greater need to apply. The possibility also exists that less stigma is attached to the receipt of transfer payments among blacks than among whites.

The final possibility is that racial discrimination does exist in the initial determination process. A racial effect is clearly present; the only question has to do with the cause of this phenomenon. All the possible causes discussed here are plausible, but no evidence exists to substantiate any of them at this time. The identification of the underlying causes of this effect must await future research. The inclusion of data on severity could alleviate this problem.