Social Security

Social Security Bulletin

Vol. 73, No. 2, 2013

IN THIS ISSUE:

The *Social Security Bulletin* (ISSN 0037-7910) is published quarterly by the Social Security Administration, 500 E Street, SW, 8th Floor, Washington, DC 20254-0001. First-class and small package carrier postage is paid in Washington, DC, and additional mailing offices.

The *Bulletin* is prepared in the Office of Retirement and Disability Policy, Office of Research, Evaluation, and Statistics. Suggestions or comments concerning the *Bulletin* should be sent to the Office of Research, Evaluation, and Statistics at the above address. Comments may also be made by e-mail at ssb@ssa.gov.

Paid subscriptions to the *Social Security Bulletin* are available from the Superintendent of Documents, U.S. Government Printing Office. The cost of a copy of the *Annual Statistical Supplement to the Social Security Bulletin* is included in the annual subscription price of the *Bulletin*. The subscription price is $56.00 domestic; $78.40 foreign. The single copy price is $13.00 domestic; $18.20 foreign. The price for single copies of the *Supplement* is $49.00 domestic; $68.60 foreign.

Internet: http://bookstore.gpo.gov
Phone: toll free (866) 512-1800;
DC area (202) 512-1800
E-mail: contactcenter@gpo.gov
Fax: (202) 512-2104
Mail: Stop IDCC, Washington, DC 20402

Postmaster: Send address changes to *Social Security Bulletin*, 500 E Street, SW, 8th Floor, Washington, DC 20254-0001.

Errata Policy: If errors that impair data interpretation are found after publication, corrections will be posted as errata on the web at http://www.socialsecurity.gov/policy/docs/ssb/v73n2/index.html.

Social Security Bulletin

Vol. 73, No. 2, 2013

Social Security Administration
Office of Retirement and Disability Policy
Office of Research, Evaluation, and Statistics

Social Security Bulletin

Volume 73 • Number 2 • 2013

Articles

Other

Subsequent Program Participation of Former Social Security Disability Insurance Beneficiaries and Supplemental Security Income Recipients Whose Eligibility Ceased Because of Medical Improvement

by Jeffrey Hemmeter and Michelle Stegman*

The Social Security Administration (SSA) periodically reviews the disabilities of Supplemental Security Income (SSI) recipients and Social Security Disability Insurance (DI) beneficiaries to determine if their impairments still meet the requirements for program eligibility. For individuals whose eligibility was ceased after a full medical review from 2003 to 2008, we track subsequent program participation for up to 8 years. We use survival analyses to estimate the time until first return to SSI and DI and explore the differences in returns by various personal and programmatic characteristics such as age, disability type, time on program, and SSA expectations regarding medical improvement. Overall, we estimate that about 30 percent of SSI-only recipients whose eligibility ceases because of medical improvement return to the SSI program within 8 years. For DI-only worker beneficiaries whose eligibility ceases, we estimate that 20 percent will return to the DI program within 8 years.

Introduction

Each year, the Social Security Administration (SSA) reviews the status of several hundred thousand Social Security Disability Insurance (DI) beneficiaries and Supplemental Security Income (SSI) recipients to determine if their medical conditions have improved enough since their last favorable determination of eligibility to allow them to engage in substantial gainful activity (SGA). To be eligible for the SSI disability program, an individual must have limited income and resources and be unable to engage in SGA because of a medically determinable physical or mental impairment that can be expected to result in death or last for at least 12 continuous months.[1] To qualify for DI, an individual must have a work history sufficient to attain insured status in addition to meeting the medical requirement.[2] At the time of award, or the last favorable review of eligibility, a date is set to revisit the individual's medical eligibility for continued participation. Because reviewing each case helps ensure that only eligible individuals receive payments, it is necessary for maintaining program integrity.

These periodic reviews, required by law, are called continuing disability reviews (CDRs). In order to keep the workload manageable and to limit administrative costs, SSA initiates the CDR process by using statistical models to identify individuals with characteristics indicating potential medical improvement. Based on those model results, SSA conducts a full medical review (FMR) only for cases deemed most likely to involve medical improvement. To individuals with a

Selected Abbreviations

CDR	continuing disability review
CIF	cumulative incidence function
DDS	Disability Determination Service
DI	Disability Insurance
FMR	full medical review

* *Jeffrey Hemmeter and Michelle Stegman are economists with the Office of Program Development, Office of Program Development and Research, Office of Retirement and Disability Policy, Social Security Administration.*

Selected Abbreviations—*Continued*

SGA	substantial gainful activity
SSA	Social Security Administration
SSI	Supplemental Security Income

lower likelihood of medical improvement, SSA sends a "mailer" asking for more information to help determine if a FMR is necessary.

During a FMR, SSA and the state Disability Determination Service (DDS) collect medical information about the participant and determine whether evidence of medical improvement exists. If the individual's condition has improved since the most recent favorable decision such that he or she is able to engage in SGA, program eligibility ceases; if not, the individual continues to receive DI benefits or SSI payments and a date is set for a future review.[3] CDRs are estimated to be highly cost effective, saving approximately $9.30 for every dollar spent on them (SSA 2012b).[4] For that reason, the 2011 deficit-reduction plan exempted CDR allocations from congressional spending caps, and the Obama administration requested an increase in CDR funding in the 2012 budget.[5]

The focus on program integrity comes at a time of substantial increases in SSI and DI participation. From 1990 through 2011, the numbers of DI beneficiaries grew from about 3.0 million to 8.6 million and disabled SSI recipients increased from 3.3 million to 6.9 million (SSA 2012c, 2012d, 2012f). Although much of the increase is simply due to the aging and growth of the population, some have argued that the programs have become relatively more attractive to low-wage individuals and those with moderate disabilities, especially during economic downturns (for example, Autor and Duggan 2003, 2006; Black, Daniel, and Sanders 2002; and Rupp and Stapleton 1998). Additionally, there is some evidence that states have transferred some of the costs formerly borne under Temporary Assistance for Needy Families onto the federal SSI program (Burkhauser and Daly 2011; Schmidt and Sevak 2004; Kubik 1999, 2003; Wamhoff and Wiseman 2005/2006). The 1990 *Sullivan v. Zebley* Supreme Court decision greatly expanded SSI eligibility for children, although welfare reform in the mid-1990s required SSA to review cases allowed during that period. Regardless, the SSI child population grew substantially in the 1990s, and many recipients continue receiving SSI as adults. In light of the increasing program costs associated with increasing participation, it is important for SSA to ensure that only those truly eligible for DI and SSI remain in the programs.

Although a few studies have looked at DI beneficiaries who medically recover (for example, Hennessey and Dykacz 1993; Dykacz and Hennessey 1989; Treitel 1979; and Schmulowitz 1973), we have not found similar studies of SSI recipients.[6] The DI studies have focused on earnings of former beneficiaries rather than subsequent program participation after a cessation decision. A few studies that look at subsequent return (Hennessey and Dykacz 1993; Dykacz and Hennessey 1989; Dykacz 1998) do not differentiate between medical and SGA-based recovery.[7]

Understanding what happens to individuals after their eligibility ceases because of medical improvement is important given recent calls across the government for stronger program integrity. Additionally, although actuaries from SSA and the Centers for Medicare and Medicaid Services incorporate returns in their models of the savings derived from CDRs, it is important for policymakers to better understand the impact of CDRs on program participation patterns.

In this article, we provide new information on the experiences of DI beneficiaries and SSI recipients after receiving a FMR that resulted in eligibility cessation. Specifically, we look at subsequent DI and SSI participation of former DI beneficiaries and former SSI recipients. Although this study does not address whether SSA's current CDR policy is adequate or how well the social safety net is working in general, we provide descriptive information on formerly eligible participants and highlight which subgroups are most likely to return to program participation.

CDR Process

The date for which a CDR is scheduled is called the CDR diary date. That date is set during the last favorable decision, which in many cases is the time of award. SSA categorizes diaries into one of three groups according to the individual's prospects for medical improvement, and the diary type determines the timing of the scheduled CDR. If medical improvement is expected, the diary date is within 3 years of the last favorable decision. For cases in which SSA deems medical improvement possible, a CDR is scheduled for 3 years after the last favorable decision. If medical improvement is not expected, a CDR is scheduled for 5 to 7 years after the last favorable decision. When the diary date approaches, SSA either "directly releases" the individual for a FMR or sends the individual a mailer containing a questionnaire seeking information to determine whether a FMR is necessary.[8]

To help determine who is directly released for a FMR and who receives a mailer, SSA uses a CDR profiling model based on administrative information to "score" the likelihood of medical improvement. SSA groups the results into three categories of likelihood of medical improvement—high, medium, or low—using cutoff scores that have not changed over time. Generally, high-scoring individuals undergo a FMR, and medium- or low-scoring DI beneficiaries and adult SSI recipients receive a mailer. However, as limited funding in recent years has restricted resources and experienced staff, SSA and the DDSs have further prioritized FMRs. As a result, some individuals do not receive their scheduled review until years later.[9]

If a mailer recipient's responses indicate medical improvement, SSA releases the case for a FMR;[10] otherwise, the agency simply sets a new CDR diary date. For a FMR, the DDS gathers medical information from the individual's medical care sources or orders consultative examinations from the treatment provider or other physicians.[11] A disability examiner and medical expert then determine if the individual's condition has improved since the last favorable decision to such an extent that he or she can perform SGA. If there has been no improvement, the individual is "continued" on the program and the DDS examiner sets a date for the next CDR. If the individual has medically improved enough to perform SGA, the examiner makes a "cessation" decision, which the individual may appeal.[12] Benefits stop after a 3-month grace period (the month of the decision and the following 2 months) unless the beneficiary appeals the decision and requests continuation of benefits during the appeal.[13] In fiscal year 2010, over 90 percent of initial CDR decisions for DI disabled-worker beneficiaries and SSI adult recipients were continuations (SSA 2012b).

The process described above has changed over time. One important example is that, as SSA moved toward statistical profiling, the agency started conducting FMRs for a sample of cases—a "profile sample"— that would not otherwise have received one. FMRs for the profile sample must be completed each year to validate the profiling model. We do not use the profile sample in our estimates because of the varying procedures under which they were drawn over the period we analyze.

Data Sources and Methodology

In this study, we use data from Social Security administrative records. The primary source is SSA's CDR Waterfall file, which contains information on all centrally initiated FMRs with a DDS determination.[14] We used an extract of the CDR Waterfall file covering calendar years 2003 through 2008.[15] That period includes FMRs conducted after the funding dedicated to processing CDRs was reduced. The file does not contain records for individuals who received a mailer unless their responses indicated possible medical improvement, in which case they went on to receive a FMR (subject to agency resources).

The file contains the date and result of the initial FMR decision by the DDS as well as the final appellate decision at the time the file was extracted. We use those data to identify records for which the FMR led to a final cessation and to define the year of the initial decision. We also use that file to create several variables likely to be correlated with return to program participation:

- CDR diary type (medical improvement expected, not expected, or possible);
- CDR profile score (high, medium, or low);
- whether the individual received a mailer or was directly released for a FMR;
- whether the individual had a prior CDR;
- whether a consultative examination was requested during the FMR;
- the adjudicative level of the decision under which the individual first entered the DI or SSI program (initial, reconsideration, Administrative Law Judge or higher, or unknown); and
- the disability considered to be the primary impairment prior to the FMR.

In addition, the file contains the date the individual became eligible for DI or SSI, the date of birth (used to establish age at the time of the initial decision[16]), sex, race, and state of residence, which may also be correlated with return to the program. For individuals receiving both SSI and DI, we use the eligibility date and adjudicative level of whichever program they entered first.

We merged the CDR Waterfall file with SSA's Numident file to obtain dates of death. If a record was missing the date of birth, we used the Supplemental Security Record and the Master Beneficiary Record (program databases covering applicants and beneficiaries for SSI and DI) to obtain it.

We also merged those files with SSA's Master Earnings File to create a measure of preeligibility earnings. We use the average earnings in the 5 years preceding the individual's date of eligibility to derive that measure. In our analyses, we include the program-specific earnings quartile of our target population. For example, we use the earnings of DI-only disabled-worker beneficiaries to define the quartiles for that group. For SSI-only recipients, we combine the two lowest quartiles because their median earnings are very close to $0.

To determine if an individual returned to DI or SSI, we merged the data described above with SSA's Disability Research File. We used that file to identify the date of the first successful postcessation application. A successful application is determined by whether benefits are awarded; postallowance technical denials are omitted. We are able to follow individuals in all of those data files through 2010.

Target Populations

All three of the target populations in this article consist of adults aged 18–59 who participated in disability programs administered by SSA until their eligibility ceased because of a FMR finding of medical improvement. The groups comprise former DI-only disabled-worker beneficiaries (individuals who did not also receive SSI payments, hereafter called "DI-only workers"), former SSI-only recipients (individuals who did not also receive DI benefits), and former disabled-worker concurrent beneficiaries (individuals who received both DI benefits and SSI payments, hereafter called "concurrent workers"). The FMRs that produced the cessation decisions were conducted during 2003–2008.

We restricted the target populations for various reasons. We removed individuals belonging to the profile sample, as well as those for whom a FMR determined reeligibility during a period of expedited reinstatement.[17] We removed records with missing or inconsistent dates, such as those indicating that an individual died before becoming eligible. We also removed individuals who appealed a cessation decision and were awaiting a new decision or still had time to file an appeal between their last cessation decision and the date the file was created. Because we focus on subsequent program participation, we excluded individuals whose eligibility did not cease. We observed the members of our sample through age 62 (as discussed below in the Analytical Methods section). Therefore, we omitted individuals who reached age 60 before their initial FMR decision and those who turned 62 before their final FMR decision in order to ensure adequate followup time.[18] We also excluded DI beneficiaries and SSI recipients who died before the final FMR decision or whose CDR profile score was missing. Those exclusions leave target populations of 33,376 DI-only workers, 24,514 SSI-only recipients, and 17,742 concurrent workers.[19] Appendix Table A-1 presents the number of records eliminated in each step of the selection process.

Limitations

SSA's CDR process is complex and dynamic. When considering our results, the reader should remember that our primary analysis pools data for several years under varying CDR policies. For example, different types of participants may have been targeted in certain years because of perceived cost savings or changes in the profiling model. Moreover, other SSA policies can also affect a CDR decision, complicating the definitional boundaries of our target populations. For example, Section 301 of the Social Security Disability Amendments of 1980 (Public Law 96-265) allows individuals to continue receiving payments even if they have received a cessation decision as long as they participate in an approved vocational program and make progress toward their employment goals. Because our observation period for each individual begins with the date of the FMR decision, the outcomes for former participants in our target population who use the Section 301 provisions and those who do not might differ. We cannot identify Section 301 use in our data (however, usage is generally low).

Our estimates also cannot anticipate future changes in funding for CDRs, the stringency of the reviews and the eligibility requirements, and the extent to which SSA uses its profiling model. The interaction of those and other factors could lead, for example, to an increase in the number of FMRs conducted. However, depending on the underlying causes and other circumstances, an increase in CDRs could result in program returns that differ in either direction from our estimates.

Analytical Methods

In this section, we discuss the cumulative incidence functions (CIFs) and proportional hazard regressions used in our analysis. We also address collinearity issues.

Cumulative Incidence Functions (CIFs)

Our primary goal is to estimate, among DI beneficiaries and SSI recipients whose eligibility ceased because of medical improvement, the percentage who subsequently returned to either the same program or the opposite program. If returning to the program was the only possible outcome and we observed all individuals over a consistent period, we would simply divide the number of individuals who returned by the number of people whose eligibility ceased. Unfortunately, neither of those conditions holds. Our observation periods range from 2 years to 8 years, depending on the year of the individual's FMR. Additionally, certain life events will compete with that outcome in other ways; death, for example, obviously precludes program return. Also, disability is no longer a factor in the SSI eligibility determination once an individual reaches age 65, and after a person reaches full retirement age (between 65 and 67 years, depending on year of birth), disability no longer affects Social Security benefit eligibility. Accurate estimates of program return must account for such factors.

To address those issues, we compute CIFs measuring the cumulative percentage of individuals from each target population who return to DI or SSI after the final cessation decision. CIFs estimate the probability of an event (such as returning to the program) when competing risks exist (Gooley and others 1999). For our analysis, we treat attainment of age 62 (which we refer to as early retirement or, simply, retirement) and death as competing events or risks.[20] Once individuals attain age 62 or die, they are no longer at risk of returning and thus provide no information about the probability of program return. Without controlling for those competing events, our estimates would assume such individuals could still return later, artificially decreasing estimated returns. Dropping the individuals who experience those events from our analysis would similarly bias our results. Thus, we estimate the probability that an individual returns to the program, allowing for the risk of dying or reaching age 62 by the end of our follow-up period (December 31, 2010). Our measure of time covers the period from the date of the final FMR decision to the first of those events.

Marubini and Valsecchi (1995) show that the CIFs can be estimated by

$$\hat{I}_j(t) = \sum_{k|t_k \le t} \hat{S}(t_k)\frac{d_{jk}}{n_k}$$

where j represents the event of interest (return to the program), $\hat{S}(t_k)$ is the overall Kaplan-Meier survival function (that is, an estimate of the probability of neither returning, dying, nor reaching age 62 by time t_k), d_{jk} is the number of individuals returning at time t_k, and n_k is the count of those at risk of returning at time t_k. Thus, it is the sum of the products of the survival estimate at time t_k and the hazard at time t_k of event j, $(\frac{d_{jk}}{n_k})$.

As described above, we are able to track program return, death, and early retirement through December 31, 2010 (the censoring date); however, we present only the results for program return. We estimate the CIFs in monthly increments and the maximum observable time span in our data is 96 months, or 8 years.[21,22]

Regressions

Because the CIF does not control for other variables that may affect return to the program, we ran Cox proportional hazard regressions on the hazard of successfully reapplying to the program to control for the characteristics of our population. Like other types of regression (such as ordinary least squares), Cox regressions provide estimates of the relative contribution of the covariates to the outcome, which in this case is the risk (or "hazard") of returning to the program over a given period of time. The exponentiated coefficients from this regression are known as hazard ratios and are interpreted similarly to odds ratios from logistic regressions: Hazard ratios greater than 1 indicate a higher risk of return relative to the reference group and those less than 1 indicate a lower risk.

The time dimension is one of the primary differences between Cox regressions and static regressions: Cox regressions estimate whether an event occurs, controlling for the timing of the event. As with the CIFs, Cox regressions control for the diverse followup times within the sample. Individuals no longer at risk of returning to the program are censored and thus drop from subsequent periods in the analysis. Unlike the CIFs, though, competing events do not hinder our ability to estimate the risk of return; that is, we can estimate the risk of return by treating competing events (death and early retirement age) as censored at the time they occur.[23]

In all our empirical models, we stratify our analyses by year of initial FMR determination, state of residence, sex, and race, allowing for separate baseline hazard functions for groups identified by those characteristics but constraining the coefficients (and hazard ratios) to be equal.[24] We do so because the different CDR policies, funding, and resources, and the variation in state policies and economies, likely affect the baseline hazard of return in each state and year in different ways. Stratification allows the effect of the other covariates in our empirical model to be proportional to the differing baseline hazards. Although this method eliminates our ability to estimate hazard ratios for the stratification variables, it also helps satisfy the proportionality assumption discussed in the following paragraph. However, future work may further consider the distributional aspects of program return.

The Cox regressions rely on the proportionality between the hazard and each covariate being constant over time. Grambsch and Therneau (1994) suggested a test of the proportionality assumption using scaled Schoenfeld residuals.[25] Those residuals (essentially the covariate value for a person actually experiencing an event minus the expected value of the event) are independent of time if the proportionality assumption is satisfied. After running that test on our empirical models, we determined that our data do not satisfy the proportionality assumption for the DI-only and concurrent worker models. For the empirical model of return to DI by former DI-only workers, the problematic variables were CDR profile score, history of a prior CDR, and preeligibility earnings quartile. In addition to those variables, the age variables did not satisfy the proportional hazards assumption in the empirical model of the return to DI by concurrent workers. For the empirical model of former DI-only workers entering SSI, the problematic variables were CDR profile score, history of a prior CDR, and mailer-recipient status. For the empirical model of concurrent workers returning to SSI, the problematic variables were history of a prior CDR and diary type.

For the problematic variables, we allow the hazard ratios to take on different values at different times. To minimize the effect of imposing a functional form on the relationship with time and to keep the empirical models computationally feasible, we allow each of the variables to have different hazard ratios for each year of followup, combining the seventh and eighth years because of small cell sizes. For example, we include a separate hazard ratio to capture the effect of a high CDR profile score in the first year after the FMR, the second year after the FMR, and so on up to 7+ years after the FMR.[26] The resulting general empirical model is:

$$h_i\left(t\right) = h_{0i}(\mathrm{t})\exp\left(\sum_{s=1}^{S}\beta_s x_s + \sum_{m=1}^{7}\gamma_m z_m\right)$$

where $h_i(t)$ is the hazard for stratification group i at time t, $h_{0i}(t)$ is the baseline hazard,[27] the β_s are the coefficients, and the x_s are the main variables. The last term on the right-hand side of the equation captures the time-varying effects, where γ_m is the effect of variable z_m m years after the FMR (and is not included in the SSI-only empirical model). In the estimation, the coefficients (β_s and γ_m) are constrained to be equal across stratification groups. All empirical models use the Efron method for treating tied events.[28]

Multicollinearity

Because of the number and the nature of the variables in our models, our estimates may suffer from multicollinearity, causing individual hazard ratios to become difficult to interpret and standard errors to be inflated. However, excluding problem variables could lead to omitted-variables bias, also causing difficult-to-interpret hazard ratios.

We tested for multicollinearity by first looking for high correlation coefficients between our variables, but did not find any we deemed especially problematic (that is, greater than 0.30). We also formally tested for multicollinearity by estimating the variance inflation factor for each variable, which is $1/(1\text{-}R^2)$ where the R^2 comes from a regression using each independent variable as the dependent variable. Because multicollinearity applies to the independent variables, functional form is irrelevant. Variance inflation factors above 10 signify multicollinearity issues. Very few variance inflation factors exceeded 4, and only one was above 10. The problematic variables were CDR profile score and years in the program. Many of our variables are included in the model estimating the CDR profile score, so its status as potentially problematic is not surprising. We also ran separate regressions subsetting on each value of our independent variables; and although hazard ratios differ across regressions, and levels of significance vary, we did not discern any consistent patterns. Additionally, there are large differences in population when we subset by those variables, which may also affect statistical significance.

Given the lack of clear evidence for multicollinearity from the variance inflation factor, low correlation coefficients, and results from subgroup-specific regressions, we do not exclude any variables from our Cox regressions or present subgroup-specific regressions. We generally focus on the direction of the hazard ratios, not their magnitudes. Thus, our regressions should be viewed as primarily exploratory or descriptive in nature, suggesting groups to focus on more closely in future research.

Characteristics of the Formerly Eligible Population

Table 1 shows demographic and programmatic characteristics of our target populations of former DI-only workers, SSI-only recipients, and concurrent workers. It covers all cases in which eligibility cessation was the outcome of a FMR conducted during calendar years 2003–2008 and for which potential appeals have expired or been exhausted. The majority (74 percent) of formerly eligible DI-only workers are aged 30–49 (with 31 percent aged 30–39 and 43 percent aged 40–49). Former SSI-only recipients are somewhat younger, with 36 percent younger than 30 and another 28 percent aged 30–39. The age distribution of concurrent workers falls somewhere in the middle, with two-thirds between ages 30 and 49 and one-quarter who are younger than 30.

The most common impairments among former DI-only workers are certain mental disorders (combined and categorized under "other mental disorders") and musculoskeletal system diseases (30 percent and 16 percent of the target population, respectively). Among former SSI-only recipients, we see the largest proportions in the other mental disorders (35 percent) and intellectual disabilities (20 percent) categories. Nearly 40 percent of former concurrent workers have other mental disorders, far outnumbering individuals in any other diagnosis category. Those impairments are similarly the most common among DI disabled-worker beneficiaries and SSI adult recipient populations overall (SSA 2012a, 2012b).

The most common diary type in each of the target populations is possible medical improvement, with 68 percent of former DI-only workers, 80 percent of former SSI-only recipients, and 65 percent of former concurrent workers. Those expected to medically improve comprise the next largest share of each target population, with 28 percent of the DI-only group, 14 percent of the SSI-only group, and 33 percent of the concurrent group. Very few individuals are not expected to medically improve. This is not surprising because those judged least likely to medically recover would generally not receive a FMR, thus excluding them from our target population.

Pluralities of former DI-only and concurrent workers (both more than 41 percent) and a majority of former SSI-only recipients (71 percent) had been program participants for 6 years or longer; another one-quarter of

Table 1.
Descriptive characteristics of former DI-only workers, SSI-only recipients, and concurrent workers whose FMRs resulted in eligibility cessation during 2003–2008

Characteristic	DI-only workers		SSI-only recipients		Concurrent workers	
	Number	Percent	Number	Percent	Number	Percent
Total	33,376	100.00	24,514	100.0	17,742	100.0
Diary type (prospective medical improvement)						
Not expected	1,226	3.7	1,468	6.0	506	2.9
Possible	22,718	68.1	19,643	80.1	11,457	64.6
Expected	9,432	28.3	3,403	13.9	5,779	32.6
CDR profile score						
Low	2,874	8.6	1,298	5.3	747	4.2
Medium	4,845	14.5	3,835	15.6	3,289	18.5
High	25,657	76.9	19,381	79.1	13,706	77.3
Age at initial CDR decision						
Younger than 30	2,926	8.8	8,729	35.6	4,442	25.0
30–39	10,189	30.5	6,920	28.2	6,057	34.1
40–49	14,348	43.0	6,787	27.7	5,718	32.2
50–59	5,913	17.7	2,078	8.5	1,525	8.6
Years in program						
Fewer than 2	974	2.9	. . .	. . .	930	5.2
2–3	9,045	27.1	. . .	. . .	4,863	27.4
Fewer than 4 (SSI only)	. . .	. . .	2,304	9.4	. . .	. . .
4–5	9,376	28.1	4,733	19.3	4,603	25.9
6 or more	13,981	41.9	17,477	71.3	7,346	41.4

(Continued)

Table 1.
Descriptive characteristics of former DI-only workers, SSI-only recipients, and concurrent workers whose FMRs resulted in eligibility cessation during 2003–2008—*Continued*

Characteristic	DI-only workers		SSI-only recipients		Concurrent workers	
	Number	Percent	Number	Percent	Number	Percent
Diagnosis						
Neoplasms	3,586	10.7	732	3.0	1,065	6.0
Intellectual disabilities	456	1.4	4,805	19.6	993	5.6
Schizophrenia and other psychotic disorders	1,984	5.9	2,381	9.7	1,711	9.6
Other mental disorders	9,916	29.7	8,603	35.1	6,976	39.3
Diseases of the—						
Endocrine, nutritional, and metabolic system	599	1.8	603	2.5	301	1.7
Nervous system and sense organs	1,852	5.6	1,378	5.6	1,026	5.8
Circulatory system	1,437	4.3	453	1.9	625	3.5
Respiratory system	531	1.6	503	2.1	302	1.7
Digestive system	1,574	4.7	381	1.6	484	2.7
Genitourinary system	1,823	5.5	540	2.2	507	2.9
Musculoskeletal system and connective tissue	5,158	15.5	1,220	5.0	1,744	9.8
Injuries	3,117	9.3	848	3.5	1,392	7.9
Other	683	2.1	413	1.7	348	2.0
Unknown [a]	660	2.0	1,654	6.8	268	1.5
Mailer receipt status						
No (direct release to FMR)	23,701	71.0	17,506	71.4	14,724	83.0
Yes	9,675	29.0	7,008	28.6	3,018	17.0
Adjudication level of initial program entry						
Initial application	720	2.2	3,683	15.0	544	3.1
Reconsideration	21,921	65.7	16,913	69.0	11,652	65.7
Administrative Law Judge or higher	3,450	10.3	2,040	8.3	2,068	11.7
Unknown	7,285	21.8	1,878	7.7	3,478	19.6
Prior CDR status						
No	26,233	78.6	18,607	75.9	13,617	76.8
Yes	7,143	21.4	5,907	24.1	4,125	23.3
Consultative examination request status						
No	19,209	57.6	12,142	49.5	9,393	52.9
Yes	14,167	42.5	12,372	50.5	8,349	47.1
Age at initial program entry (SSI only)						
Younger than 18	. . .	. . .	6,991	28.5	. . .	. . .
18 or older	. . .	. . .	17,523	71.5	. . .	. . .
Calendar year of FMR						
2003	7,582	22.7	9,888	40.3	4,022	22.7
2004	7,640	22.9	8,110	33.1	4,076	23.0
2005	8,066	24.2	4,295	17.5	4,162	23.5
2006	4,782	14.3	1,130	4.6	2,677	15.1
2007	3,084	9.2	678	2.8	1,752	9.9
2008	2,222	6.7	413	1.7	1,053	5.9

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Rounded components of percentage distributions do not necessarily sum to 100.

. . . = not applicable.

a. Impairment type missing from CDR Waterfall data file.

DI-only and concurrent workers had participated for 4 to 5 years. Those large shares may result from a decline in CDR funding and a growing backlog of cases. We note that over three-quarters of each target population did not have a CDR prior to the current one, and over 60 percent have had medical improvement deemed possible (meaning a CDR scheduled every 3 years).

About 70 percent of DI-only and concurrent workers and 90 percent of SSI-only recipients had their FMR during the first half of our study period (2003–2005). The decline in FMRs in the latter half of the period is most likely due to a decrease in the number of cases sent for review because of lower funding. Year-to-year differences in the percentage of FMRs may also be related to changing CDR policies in SSA. Appendix Tables B-1 through B-3 report statistics for each target population in the first (2003) and last (2008) FMR years we analyze.[29]

Return to DI and SSI

In this section, we present estimates of the return to DI and SSI within 8 years of a final cessation decision. We begin with the estimates of the CIFs for the full target populations and follow with estimates for subsetting characteristics. We then turn to the regression results, focusing separately on each target population's return to DI and SSI.

CIF Results

We estimate the CIFs of return to DI and SSI for each beneficiary type, that is, the probability that a former participant successfully applies for DI or SSI by a given month. As stated earlier, we follow individuals until they successfully reapply for SSI or DI (depending on the empirical model), they attain age 62, they die, or December 31, 2010, whichever occurred first.[30] We present estimates of program return to DI in Chart 1 and to SSI in Chart 2.

Recall that we are measuring the time from the final FMR cessation decision to the application that leads to a new award. Given the large volume of appeals and the SSA backlog, it likely takes several more months until the first payment is received by those who return. However, in most circumstances, back payments will cover the time from favorable eligibility determination to first payment.

We estimate that about 20 percent of our DI-only target population and 21 percent of concurrent workers will return to DI within 8 years of an eligibility cessation due to medical improvement (Chart 1). More than one-half of those returns occur within the first few years of the FMR—at 3 years, roughly 11 percent of each group had returned.

A much smaller percentage of the SSI-only group successfully applies for DI after their SSI eligibility ceases (6 percent). Former SSI-only recipients must establish a sufficient work history to become eligible for DI. We cannot determine how many quarters of coverage those individuals had prior to entering SSI; some may only have needed a few quarters while others may have needed many. However, former SSI-only recipients with higher preparticipation earnings are more likely to subsequently enter DI than those with lower preparticipation earnings.

We estimate that almost 30 percent of the SSI-only group will return to SSI within 8 years of a final eligibility cessation (Chart 2). Unsurprisingly, concurrent workers return to SSI at about the same rate as they return to DI (22 percent). We estimate that 11 percent of former DI-only workers will successfully apply for SSI payments within 8 years.[31]

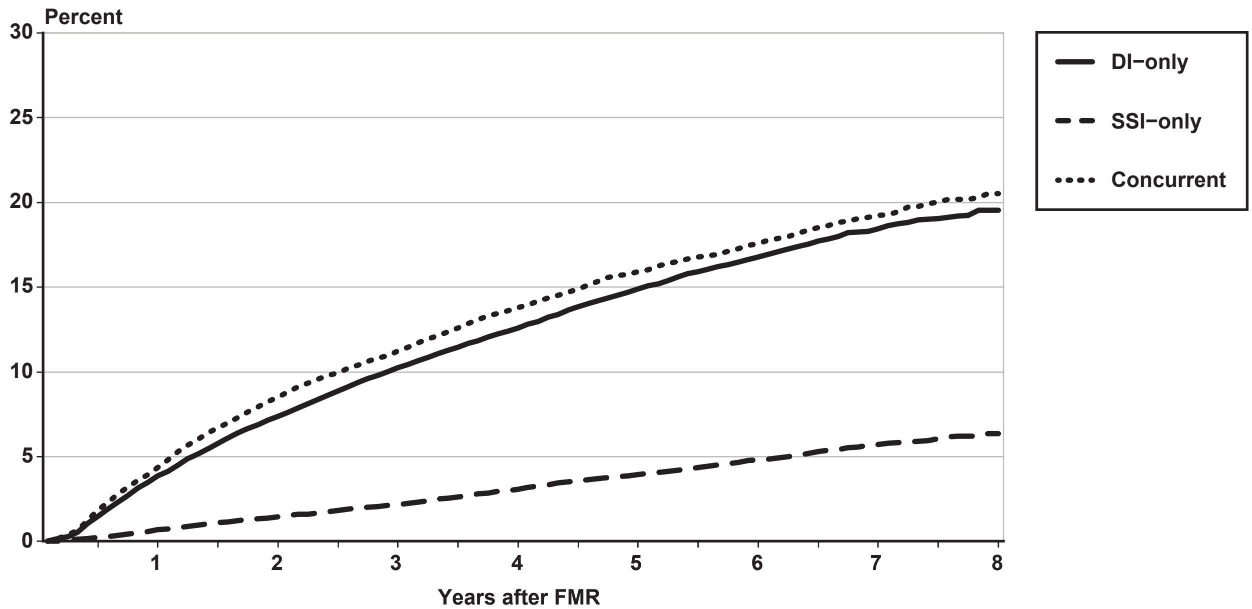
Note that the estimated CIFs at year 8 reflect the experiences of the earliest FMRs in our target population. However, the greatest risk of return, measured by the slope of the CIF, is in the first few years after the FMR. Although CIFs increase over time, they do so at diminishing rates.[32]

CIFs by Subsetting Characteristics

Table 2 presents the estimated cumulative incidence of successfully applying for DI or SSI within 8 years of cessation for each target population by characteristic. The first line replicates the final values of the overall CIFs in Charts 1 and 2 (that is, the average return after 8 years).

The estimated percentages of successful DI or SSI application vary substantially across characteristics. In general, those for whom SSA does not expect medical improvement are more likely to return within 8 years than the groups for whom medical improvement is expected or deemed possible. A higher percentage of older individuals tend to return to their original program (or to either program for former concurrent workers), compared with the overall return averages. The return percentage for those with a prior CDR is lower than average across all categories; correspondingly, the percentage is higher than average among those without a prior CDR.

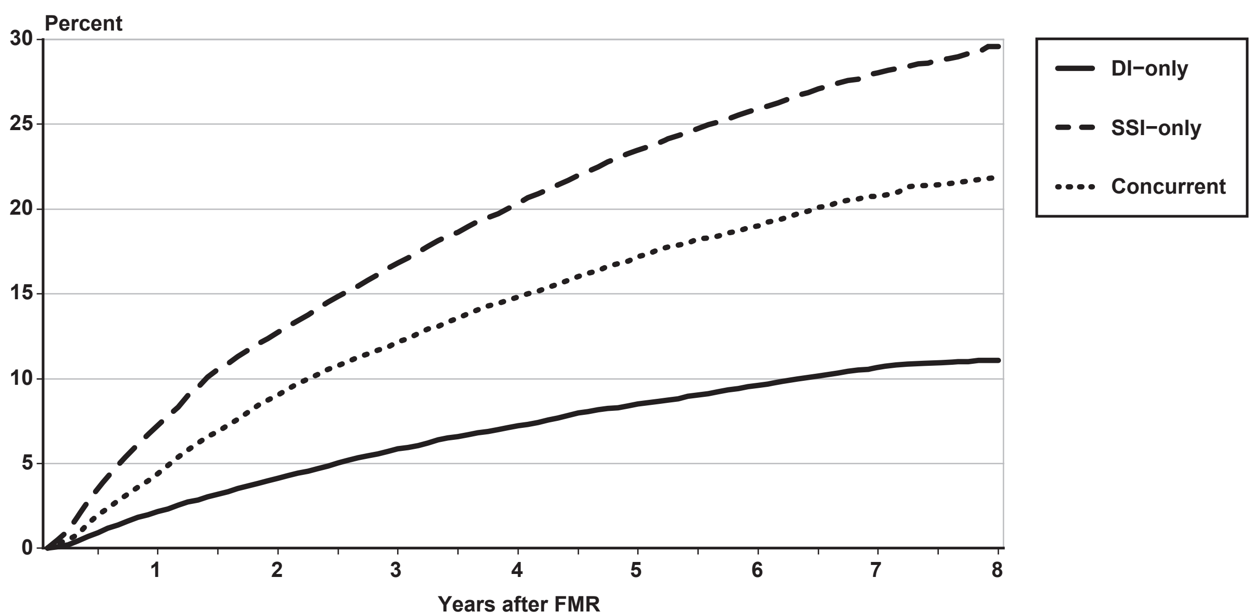
Chart 1.
Estimated percentage of former DI-only workers, SSI-only recipients, and concurrent workers who successfully applied to DI after their FMR cessation decision



SOURCE: Authors' calculations using Social Security administrative records.

NOTE: Covers cases with cessation decisions reached in FMRs conducted in 2003–2008, and followed through 2010.

Chart 2.
Estimated percentage of former DI-only workers, SSI-only recipients, and concurrent workers who successfully applied to SSI after their FMR cessation decision



SOURCE: Authors' calculations using Social Security administrative records.

NOTE: Covers cases with cessation decisions reached in FMRs conducted in 2003–2008, and followed through 2010.

Table 2.
Cumulative incidence of successful reapplication to DI or SSI after a FMR cessation decision reached during 2003–2008, by former program type and beneficiary characteristics (in percent)

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
Total	19.54	11.07	6.37	29.59	20.52	21.79
Diary type (prospective medical improvement)						
Not expected	26.77	17.14	8.57	31.09	27.25	28.45
Possible	18.96	11.48	6.27	29.78	19.64	21.51
Expected	19.99	9.39	6.10	27.74	21.66	20.92
CDR profile score						
Low	17.76	10.96	4.07	32.26	22.30	27.02
Medium	19.30	10.30	5.98	33.60	21.33	22.77
High	19.80	11.22	6.62	28.69	20.28	20.91
Age at initial CDR decision						
Younger than 30	16.02	11.41	6.53	22.05	18.12	17.98
30–39	15.90	10.62	6.31	28.65	17.96	19.75
40–49	21.82	11.84	6.33	37.84	22.72	24.49
50–59	22.37	9.93	5.94	38.83	30.29	28.48
Years in program						
Fewer than 2	22.52	6.33	. . .	. . .	20.69	16.99
2–3	22.10	10.80	. . .	. . .	23.61	21.72
Fewer than 4 (SSI only)	. . .	. . .	6.00	30.53	. . .	. . .
4–5	22.71	12.99	6.35	35.80	22.39	25.51
6 or more	15.22	10.32	6.43	27.73	16.82	19.40
Diagnosis						
Neoplasms	18.53	7.14	6.33	18.42	19.68	16.91
Intellectual disabilities	23.27	22.37	6.31	26.95	20.34	27.20
Schizophrenia and other psychotic disorders	28.37	22.17	6.56	38.46	25.32	28.30
Other mental disorders	18.28	11.13	5.71	27.71	20.16	20.78
Diseases of the—						
Endocrine, nutritional, and metabolic system	22.60	11.22	9.88	34.74	22.51	19.90
Nervous system and sense organs	16.64	10.71	7.15	30.24	17.77	19.24
Circulatory system	27.01	13.79	10.20	41.20	27.25	26.05
Respiratory system	24.09	14.32	3.72	27.33	19.70	22.67
Digestive system	18.34	9.19	3.51	29.40	18.86	16.79
Genitourinary system	30.42	13.04	12.88	34.55	26.04	23.49
Musculoskeletal system and connective tissue	17.11	9.08	5.31	33.51	20.61	22.16
Injuries	15.07	8.98	4.93	25.72	15.59	17.08
Other	19.51	11.22	9.48	27.56	22.10	21.78
Unknown [a]	14.28	8.31	6.18	31.23	15.96	20.23
Mailer receipt status						
No (direct release to FMR)	19.91	10.97	6.70	28.95	20.21	21.11
Yes	18.52	11.22	5.69	31.28	22.31	23.85
Adjudication level of initial program entry						
Initial application	20.62	10.99	6.42	28.89	20.99	21.61
Reconsideration	19.13	11.33	6.40	31.28	21.44	22.52
Administrative Law Judge or higher	16.39	10.76	5.45	32.67	18.42	20.93
Unknown	19.27	14.18	6.62	30.41	19.34	19.92
Prior CDR status						
No	22.68	12.59	6.48	33.06	23.45	24.48
Yes	7.31	5.19	6.02	18.19	10.71	11.28

(Continued)

Table 2.
Cumulative incidence of successful reapplication to DI or SSI after a FMR cessation decision reached during 2003–2008, by former program type and beneficiary characteristics (in percent)—*Continued*

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
Consultative examination request status						
No	19.86	11.30	6.91	31.08	21.33	22.48
Yes	19.06	10.70	5.81	28.13	19.60	20.37
Age at initial program entry (SSI only)						
Younger than 18	. . .	. . .	5.99	21.39	. . .	. . .
18 or older	. . .	. . .	6.52	32.85	. . .	. . .
Preeligibility earnings quartile						
Lowest	17.97	15.36	. . .	. . .	16.26	21.08
Second	21.33	13.87	. . .	. . .	19.69	23.73
Lowest or second (SSI only)	. . .	. . .	5.47	28.03	. . .	. . .
Third	20.20	10.17	6.64	30.91	22.16	21.45
Highest	18.74	4.99	7.87	31.48	23.96	19.91

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Covers cases with cessation decisions reached in FMRs conducted in 2003–2008, and followed through 2010.

. . . = not applicable.

a. Impairment type missing from CDR Waterfall data file.

Among those diagnosed with schizophrenia and other psychotic disorders, circulatory system diseases, and genitourinary system diseases, we estimate higher-than-average percentages returning to each program from all three former program types. We estimate lower-than-average percentages returning among those with neoplasms, digestive system diseases, and injuries from all three target populations. The other characteristic groups show less variation across program types.

Regression Results

We estimated Cox proportional hazard regressions of the time to first successful postcessation DI or SSI application, controlling for the characteristics described earlier. Table 3 presents the hazard model results for program returns. The aggregate hazard ratios for the entire study period appear in the upper panel of Table 3 and the hazard ratios of the time-varying effects in each model are shown in the lower panel. Recall that the variables we include as time-varying are those that did not satisfy the proportionality assumptions of each Cox regression. Note that the methodology we use to estimate the time-varying effects creates separate observations for each distinct time period during which we observe an individual. Thus, an individual who, for example, has a medium CDR profile score and is observed for 4 years in the DI-only regressions will have four different observations in the data, one for each calendar year after cessation. As a result, the number of observations for DI-only and concurrent regressions shown in Table 3 is substantially higher than the total sample values given in Table 1; but the observations for the SSI-only regressions, which do not include time-varying effects, match the Table 1 values. Appendix Table C-1 presents standard errors for the regressions.

Former DI-only Workers

All else being equal, former DI-only workers have a higher risk (hazard) of returning to DI if they were older or judged less likely to improve according to the diary type. To illustrate, the hazard ratio of 1.40 for the medical improvement not expected group implies that the group, in any given year after cessation, had 1.40 times the risk of returning to DI as did the reference group (for which medical improvement was expected). Alternatively, those with higher CDR profile scores (that is, more likely to have their eligibility ceased according to SSA's profiling model) have a lower risk of return—although this effect diminishes after 3 years. For example, in the first year after cessation, the high CDR profile-score group's hazard ratio of 0.73 indicates that the risk of return to DI for

Table 3.
Proportional hazard regression results (hazard ratios) of time to first successful application to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Aggregate effects			
Diary type (prospective medical improvement)						
Not expected	1.40***	1.43***	2.26***	1.09	1.31**	a
Possible	1.14***	1.22***	1.21*	1.07	1.10	a
Expected (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
CDR profile score						
Low (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Medium	a	a	0.88	0.91	a	0.83*
High	a	a	1.33	0.97	a	0.78**
Age at initial CDR decision						
Younger than 30 (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
30–39	1.25***	1.35***	0.80**	1.27***	a	1.37***
40–49	1.83***	1.75***	0.87	1.83***	a	1.91***
50–59	2.32***	1.76***	1.22	2.20***	a	2.37***
Years in program						
Fewer than 2 (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .	. . .	. . .
2–3	0.94	1.18	. . .	. . .	1.03	1.12
Fewer than 4 (reference group for SSI-only)	. . .	. . .	. . .	. . .	. . .	. . .
4–5	0.96	1.21	0.78*	1.09	0.97	1.16
6 or more	0.76***	1.08	0.85*	1.19***	0.86	1.02
Diagnosis						
Neoplasms	0.98	0.86	0.74	0.54***	0.82*	0.80**
Intellectual disabilities	1.34**	1.71***	1.22	1.21***	1.13	1.26**
Schizophrenia and other psychotic disorders	1.92***	2.17***	1.30	1.41***	1.56***	1.44***
Other mental disorders	1.17***	1.34***	1.01	1.06	1.06	1.11
Diseases of the—						
Endocrine, nutritional, and metabolic system	1.20*	1.13	1.58*	1.27**	1.11	0.91
Nervous system and sense organs	1.02	1.13	1.21	1.06	0.96	0.99
Circulatory system	1.35***	1.49***	1.18	1.15	1.14	1.16
Respiratory system	1.13	1.15	0.74	1.02	1.01	1.05
Digestive system	0.92	0.99	0.72	0.84	0.75**	0.81
Genitourinary system	1.48***	1.41***	2.16***	1.21*	1.32**	1.15
Musculoskeletal system and connective tissue (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Injuries	0.83***	0.96	0.92	0.93	0.69***	0.79**
Other	1.09	1.18	1.59*	1.12	1.14	1.23
Unknown [b]	0.87	0.86	1.16	1.01	0.77	0.85
Mailer receipt status						
No (direct release to FMR; reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes	0.95	a	0.57***	0.99	1.02	1.00
Adjudication level of initial program entry						
Initial application (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Reconsideration	0.96	0.97	0.82	0.94	1.10	1.09
Administrative Law Judge or higher	0.88***	0.99	0.80*	1.01	0.91	0.96
Unknown	1.04	1.19	1.08	0.99	0.98	0.92

(Continued)

Table 3.
Proportional hazard regression results (hazard ratios) of time to first successful application to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Aggregate effects (cont.)			
Prior CDR status						
No (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes	a	a	1.00	0.58***	a	a
Consultative examination request status						
No (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes	0.94**	0.93*	0.89*	0.87***	0.84***	0.84***
Age at initial program entry (SSI only)						
Younger than 18 (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
18 or older	. . .	. . .	1.16	1.09*	. . .	. . .
Preeligibility earnings quartile						
Lowest (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .	. . .	. . .
Second	a	0.89**	. . .	. . .	a	1.08
Lowest or second (reference group for SSI-only)	. . .	. . .	. . .	. . .	. . .	. . .
Third	a	0.62***	1.20**	0.93*	a	0.90*
Highest	a	0.30***	1.63***	0.82***	a	0.73***
			Time-varying effects			
Diary type (prospective medical improvement)						
Not expected						
Year 1	c	c	c	c	c	1.35
Year 2	c	c	c	c	c	1.57**
Year 3	c	c	c	c	c	0.90
Year 4	c	c	c	c	c	1.06
Year 5	c	c	c	c	c	1.71*
Year 6	c	c	c	c	c	0.42
Year 7 or 8	c	c	c	c	c	0.35
Possible						
Year 1	c	c	c	c	c	1.23**
Year 2	c	c	c	c	c	1.06
Year 3	c	c	c	c	c	1.09
Year 4	c	c	c	c	c	1.00
Year 5	c	c	c	c	c	1.09
Year 6	c	c	c	c	c	1.18
Year 7 or 8	c	c	c	c	c	0.92
Expected (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
CDR profile score						
Low (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Medium						
Year 1	0.81*	0.62***	c	c	1.70**	c
Year 2	0.61***	0.72*	c	c	0.76	c
Year 3	0.74**	0.70*	c	c	0.65*	c
Year 4	1.16	0.82	c	c	1.79	c
Year 5	1.28	1.36	c	c	0.86	c
Year 6	1.13	0.63	c	c	0.23***	c
Year 7 or 8	0.78	1.76	c	c	1.60	c

(Continued)

Table 3.
Proportional hazard regression results (hazard ratios) of time to first successful application to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Time-varying effects (cont.)			
CDR profile score (cont.)						
High						
Year 1	0.73***	0.77	c	c	1.15	c
Year 2	0.68***	0.72*	c	c	0.75	c
Year 3	0.74**	0.72*	c	c	0.68	c
Year 4	1.06	0.62**	c	c	1.97*	c
Year 5	0.99	1.60	c	c	0.82	c
Year 6	1.25	0.49**	c	c	0.42**	c
Year 7 or 8	0.75	1.41	c	c	0.99	c
Age at initial CDR decision						
Younger than 30 (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
30–39						
Year 1	c	c	c	c	1.47***	c
Year 2	c	c	c	c	1.35***	c
Year 3	c	c	c	c	0.99	c
Year 4	c	c	c	c	1.16	c
Year 5	c	c	c	c	1.25	c
Year 6	c	c	c	c	1.17	c
Year 7 or 8	c	c	c	c	0.73	c
40–49						
Year 1	c	c	c	c	1.80***	c
Year 2	c	c	c	c	1.79***	c
Year 3	c	c	c	c	1.54***	c
Year 4	c	c	c	c	1.50**	c
Year 5	c	c	c	c	1.72***	c
Year 6	c	c	c	c	1.94**	c
Year 7 or 8	c	c	c	c	0.76	c
50–59						
Year 1	c	c	c	c	2.61***	c
Year 2	c	c	c	c	2.51***	c
Year 3	c	c	c	c	2.21***	c
Year 4	c	c	c	c	2.81***	c
Year 5	c	c	c	c	2.85***	c
Year 6	c	c	c	c	2.77**	c
Year 7 or 8	c	c	c	c	0.44	c
Mailer receipt status						
No (direct release to FMR; reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes						
Year 1	c	1.11	c	c	c	c
Year 2	c	0.98	c	c	c	c
Year 3	c	1.15	c	c	c	c
Year 4	c	0.59***	c	c	c	c
Year 5	c	1.08	c	c	c	c
Year 6	c	0.78	c	c	c	c
Year 7 or 8	c	0.90	c	c	c	c

(Continued)

Table 3.
Proportional hazard regression results (hazard ratios) of time to first successful application to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Time-varying effects (cont.)			
Prior CDR status						
No (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes						
Year 1	0.21***	0.17***	c	c	0.19***	0.23***
Year 2	0.31***	0.29***	c	c	0.24***	0.28***
Year 3	0.28***	0.34***	c	c	0.37***	0.40***
Year 4	0.35***	0.42***	c	c	0.35***	0.37***
Year 5	0.34***	0.42***	c	c	0.42***	0.37***
Year 6	0.45***	0.50***	c	c	0.48***	0.42***
Year 7 or 8	0.27***	0.25***	c	c	0.43***	0.41***
Preeligibility earnings quartile						
Lowest (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .	. . .	. . .
Second						
Year 1	1.10	c	. . .	. . .	1.19	c
Year 2	1.14	c	. . .	. . .	1.13	c
Year 3	0.97	c	. . .	. . .	1.29*	c
Year 4	1.05	c	. . .	. . .	1.41*	c
Year 5	1.25*	c	. . .	. . .	1.42*	c
Year 6	2.06***	c	. . .	. . .	1.58*	c
Year 7 or 8	1.15	c	. . .	. . .	1.21	c
Third						
Year 1	1.00	c	c	c	1.21	c
Year 2	0.92	c	c	c	1.05	c
Year 3	0.82*	c	c	c	1.48**	c
Year 4	1.01	c	c	c	1.49**	c
Year 5	0.96	c	c	c	1.53**	c
Year 6	1.43*	c	c	c	1.35	c
Year 7 or 8	1.56**	c	c	c	1.75*	c
Highest						
Year 1	0.75***	c	c	c	1.21	c
Year 2	0.78***	c	c	c	0.95	c
Year 3	0.74***	c	c	c	1.19	c
Year 4	0.70***	c	c	c	1.83***	c
Year 5	0.96	c	c	c	1.30	c
Year 6	1.52**	c	c	c	1.00	c
Year 7 or 8	0.83	c	c	c	2.60***	c
Observations	168,675	174,736	24,514	24,514	87,471	87,050

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Covers cases with cessation decisions reached in FMRs conducted in 2003–2008, and followed through 2010.

. . . = not applicable.

* = statistically significant at the 0.1 level.

** = statistically significant at the 0.05 level.

*** = statistically significant at the 0.01 level.

a. Included as a time-varying effect because the CIF did not satisfy the proportionality assumption. See lower panel.

b. Impairment type missing from CDR Waterfall data file.

c. No time-varying Cox regression was calculated because the CIF (shown in the upper panel) satisfied the proportionality assumption.

members of this group was only 73 percent of that for members of the low profile-score group. In the fourth year after cessation, however, there is no difference in risk of return between the two groups (the hazard ratio is 1.06 and is not statistically significant). Former DI-only workers with a lower risk of return include those who had a prior CDR, those who required a consultative examination, and those who were on the DI program for 6 or more years (compared with those who were on DI for fewer than 2 years).

Former DI-only workers in the highest preeligibility earnings quartiles are less likely to return to DI within 4 years than are those in the lowest quartile, all else being equal. Relative to those with musculoskeletal system and connective tissue impairments, individuals with intellectual disabilities are much more likely to return to DI, as are those with schizophrenia and other psychotic disorders; other mental disorders; endocrine, nutritional, and metabolic diseases; circulatory system diseases; and genitourinary system diseases, all else being equal. Individuals with injuries have a lower risk of return than do those with musculoskeletal impairments. Also, those initially allowed at the Administrative Law Judge level or higher have a lower risk of return to DI than do those allowed at the initial adjudication level.

Although the magnitudes differ, the signs and significance of the hazard ratios of subsequent SSI participation for former DI-only workers are generally similar to those for subsequent DI participation. The hazard ratios of individuals previously on DI for 6 or more years, those allowed at the Administrative Law Judge level or higher, and those with injuries are not significant in the SSI empirical model. Consistent with the means-tested nature of SSI, former DI-only workers in higher preeligibility earnings quartiles have a lower risk of successfully applying for SSI than do those with earnings in the lowest quartile.

Former SSI-only Recipients

All else held equal, former SSI-only recipients have a higher risk of successfully applying for DI if they are considered less likely to medically improve (as judged by diary type) and if they had higher preeligibility earnings. Former SSI-only recipients who were on the program for 4 years or more, received a mailer, or required a consultative examination have a lower risk of successfully applying for DI. Additionally, those with endocrine, nutritional, and metabolic diseases, genitourinary system diseases, and "other" impairments are more likely than those with musculoskeletal and connective tissue impairments to apply successfully for DI.

The characteristics influencing return to SSI by former SSI-only recipients differ from those influencing successful application for DI. For example, the diary type and mailer status hazard ratios are not statistically significant in the SSI regression. Additionally, those with a prior CDR are less likely to return to SSI, and those who were aged 18 or older at the time they first entered SSI are more likely to return to SSI. Neither of those variables is significant in the DI-return model. Older individuals are also more likely to return to SSI. As would be expected, those with higher preeligibility earnings are less likely to return to SSI, although we found them more likely to successfully apply for DI after SSI cessation.

Former Concurrent Workers

In the empirical models for former concurrent workers, those not expected to medically improve are more likely to return to each program, but those with medical improvement deemed possible are more likely to return only to SSI. In the SSI empirical model, those effects are sporadic; in cases where medical improvement is not expected, the hazard ratios are statistically significant in only the second and fifth years (1.57 and 1.71, respectively), and where improvement is deemed possible, only the first-year estimate (1.23) is significant. Individuals with higher CDR profile scores are less likely to return to SSI, but those effects fluctuate in the DI empirical model, with some hazard ratios above 1 and others below 1 in no consistent pattern. In both empirical models, those with a prior CDR and those who required a consultative examination are less likely to return to DI and SSI. The hazard ratios for the highest two earnings quartiles in the SSI-return empirical model are statistically significant, with individuals in those quartiles less likely to return to SSI. In the DI empirical model, the estimates suggest higher earners are somewhat more likely to enter DI, but the hazard ratios vary over the followup period. Older individuals are also more likely to return to each program.

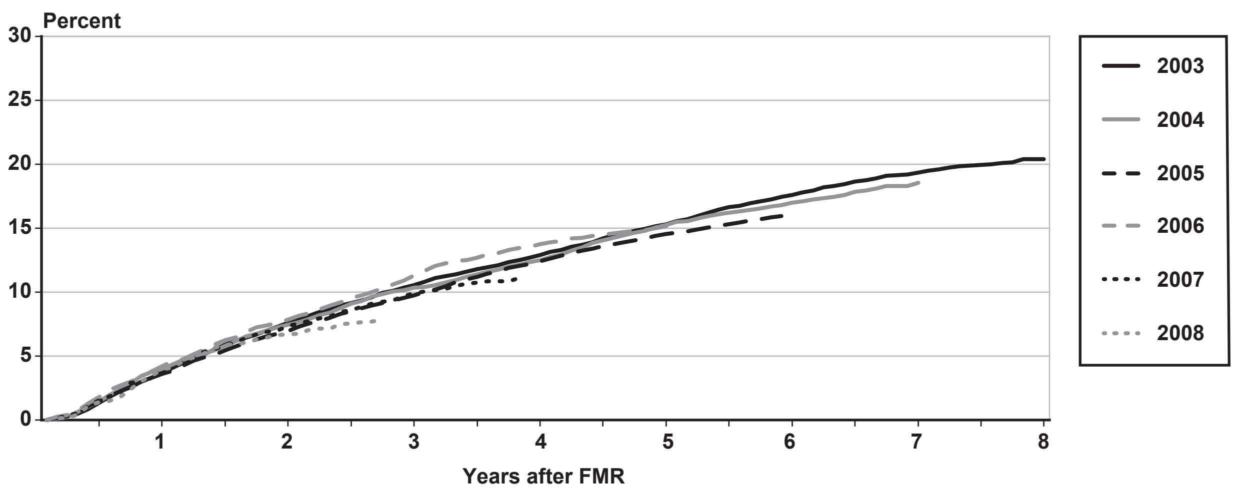
Individuals with schizophrenia and other psychotic disorders are more likely to return to either program than are those with musculoskeletal and connective tissue impairments; those with neoplasms and injuries are less likely to return. Former concurrent workers with intellectual disabilities are more likely to return to the SSI program. As for the DI program, individuals with digestive systems diseases are less likely to return, while those with genitourinary system diseases are more likely.

Year-Specific Estimates

As discussed earlier, our aggregate results pool several cohort years together, resulting in heterogeneous target populations. Therefore, the estimated CIFs may mask differences in the rates of program return between yearly cohorts. To explore that possibility, we present the estimates of the CIFs for each FMR cohort year for former DI-only workers returning to DI (Chart 3) and former SSI-only recipients returning to SSI (Chart 4) through the maximum followup time.[33]

For both programs, there is substantial overlap of the cohort-year estimates over time—program return is fairly similar in every followup month for each yearly cohort. However, for former DI-only workers, there is some evidence of a downward shift—the curves are somewhat flatter in successive cohorts. We compared the 95-percent confidence intervals of

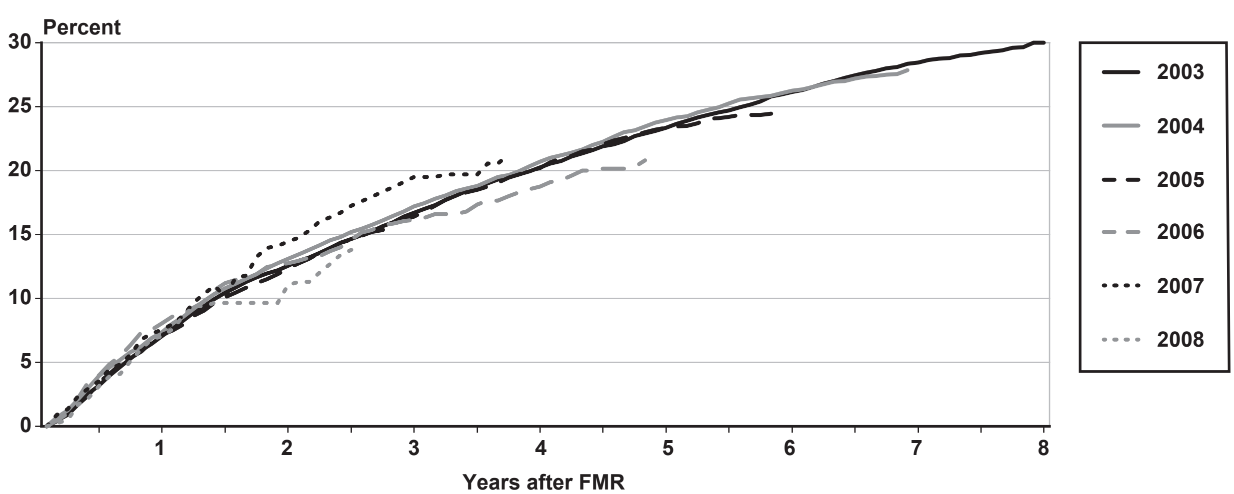
Chart 3.
Estimated percentage of former DI-only workers who successfully reapplied to DI after their FMR cessation decision, by FMR year



SOURCE: Authors' calculations using Social Security administrative records.

NOTE: Covers cases with cessation decisions reached in FMRs conducted in 2003–2008, and followed through 2010.

Chart 4.
Estimated percentage of former SSI-only recipients who successfully reapplied to SSI after their FMR cessation decision, by FMR year



SOURCE: Authors' calculations using Social Security administrative records.

NOTE: Covers cases with cessation decisions reached in FMRs conducted in 2003–2008, and followed through 2010.

the 2003 and 2008 cohorts, the earliest and latest in our sample, to determine the extent of that trend. The confidence intervals for those two cohorts overlap for all but the last 3 months of their common followup time (not shown in the charts). The difference between those two cohorts at the end of the common followup period is about 2 percentage points, but over the first year and a half they are virtually identical.[34]

That finding may result from a tightening of CDR funding over the period—inflation-adjusted CDR funding decreased from about $659 million in fiscal year 2003 to just over $300 million in fiscal year 2008.[35] With the drop in funding, SSA reduced the number of FMRs (for both SSI and DI) by about 400,000. Combined with the improved profiling models used during the period, the fewer FMRs were increasingly targeted to individuals less likely to qualify for benefits and arguably less likely to return to the program. Following the later cohorts for longer periods will help determine whether this is a long-standing result or an inconsequential blip in the data.[36]

Based on a comparison of the confidence intervals, a similar trend does not appear among former SSI-only recipients, which may be due to the smaller populations with ceased SSI eligibility in each year (down to just over 400 in 2008; the confidence intervals overlap for all years). Plots of cross-program participation and former concurrent beneficiary returns show trends similar to those for same-program returns (not shown).

For the Cox regressions, recall that stratification imposes identical hazard ratio estimates on each yearly stratum. To obtain yearly estimates, we also ran proportional hazard regressions for each yearly cohort to reveal any systematic changes in the estimated hazard ratios over time. Table 4 presents year-specific Cox regressions of same-program return (Appendix Table D-1 presents standard errors). For the DI-only population, we show regressions for the 2003 and 2008 cohorts. For the SSI-only population we show regressions for the 2003 cohort and, because the 2008 cohort is small, a pooled 2007/2008 cohort. We limit the regressions to the maximum followup period for the 2008 cohort (36 months, counting the month of eligibility cessation as month 1). As in the prior regressions, we continue to stratify by state, sex, and race, and allow for time-varying effects of variables that do not pass proportional hazards tests. Additionally, some variable categories needed to be combined because of small sample sizes; thus, the yearly models differ from those for pooled regressions.

Some hazard ratios change in magnitude and for others the direction of the risk of return changes. The only effect that is statistically significant and consistent across target populations for both years is the decreased risk of returning for those who have had a prior CDR. We also see an increased risk of return for individuals who are older (with the exception of the 2007/2008 SSI regression). In general, few

Table 4.
Proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 3 years of a 2003 or 2008 FMR cessation decision, by selected beneficiary characteristics

Characteristic	Former DI-only workers, returned to DI within 3 years of FMR in—		Former SSI-only recipients, returned to SSI within 3 years of FMR in—	
	2003	2008	2003	2007/2008 [a]
	Aggregate effects			
Diary type (prospective medical improvement)				
Not expected	0.91	1.50	1.00	0.97
Possible	1.00	1.04	1.11*	0.99
Expected (reference group)	. . .	. . .	. . .	. . .
CDR profile score				
Low (reference group)	. . .	. . .	. . .	. . .
Medium	0.93	0.62	1.03	0.87
High	b	b	1.06	0.79
Age at initial CDR decision				
Younger than 30 (reference group)	. . .	. . .	. . .	. . .
30–39	b	b	1.28***	0.67
40–49	1.68***	2.11*	1.81***	0.74
50–59	2.13***	3.27***	2.01***	1.43

(Continued)

Table 4.
Proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 3 years of a 2003 or 2008 FMR cessation decision, by selected beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, returned to DI within 3 years of FMR in—		Former SSI-only recipients, returned to SSI within 3 years of FMR in—	
	2003	2008	2003	2007/2008 [a]
	Aggregate effects (cont.)			
Years in program				
Fewer than 4 (reference group for DI-only)	. . .	. . .	. . .	. . .
4–5	0.97	1.13	. . .	. . .
Fewer than 6 (reference group for SSI-only)	. . .	. . .	. . .	. . .
6 or more	0.95	0.77	0.88***	0.77
Diagnosis				
Neoplasms	0.97	0.69	c	c
Intellectual disabilities	1.58*	2.71	1.22***	1.15
Schizophrenia and other psychotic disorders	1.86***	1.08	[d] 1.17***	[d] 0.91
Other mental disorders	1.16	0.86	[d] 1.17***	[d] 0.91
Diseases of the—				
Endocrine, nutritional, and metabolic system	b	b	c	c
Nervous system and sense organs	0.97	0.66	0.92	0.84
Circulatory system	1.46***	1.58	c	c
Respiratory system	1.03	1.55	c	c
Digestive system	0.80	0.38	c	c
Genitourinary system	b	b	c	c
Musculoskeletal system and connective tissue (reference group)	. . .	. . .	. . .	. . .
Injuries	0.77**	0.94	c	c
Other	0.88	0.74	c	c
Unknown [e]	0.88	1.55	c	c
Mailer receipt status				
No (direct release to FMR; reference group)	. . .	. . .	. . .	. . .
Yes	1.08	0.75	1.11*	0.99
Adjudication level of initial program entry				
Initial application (reference group)	. . .	. . .	. . .	. . .
Reconsideration	1.02	0.62	0.97	0.76
Administrative Law Judge or higher	0.92	1.09	1.06	0.59
Unknown	0.90	1.52	0.94	0.91
Prior CDR status				
No (reference group)	. . .	. . .	. . .	. . .
Yes	0.37***	0.12***	0.65***	0.48***
Consultative examination request status				
No (reference group)	. . .	. . .	. . .	. . .
Yes	0.98	0.92	0.89***	0.80
Age at initial program entry (SSI only)				
Younger than 18 (reference group)	. . .	. . .	. . .	. . .
18 or older	. . .	. . .	1.04	1.05
Preeligibility earnings quartile				
Lowest (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .
Second	1.27***	1.00	. . .	. . .
Lowest or second (reference group for SSI-only)	. . .	. . .	. . .	. . .
Third	1.07	0.84	0.98	1.11
Highest	0.87	0.78	0.86***	0.95

(Continued)

Table 4.
Proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 3 years of a 2003 or 2008 FMR cessation decision, by selected beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, returned to DI within 3 years of FMR in— 2003	2008	Former SSI-only recipients, returned to SSI within 3 years of FMR in— 2003	2007/2008 [a]
	Time-varying effects			
CDR profile score				
Low (reference group)	. . .	. . .	. . .	. . .
High				
Year 1	0.68*	0.70	f	f
Year 2	1.07	0.41*	f	f
Year 3	1.17	0.14***	f	f
Age at initial CDR decision				
Younger than 30 (reference group)	. . .	. . .	. . .	. . .
30–39				
Year 1	1.05	0.87	f	f
Year 2	1.28	1.39	f	f
Year 3	1.10	2.54	f	f
Diagnosis				
Diseases of the—				
Endocrine, nutritional, and metabolic system				
Year 1	0.28*	2.84	f	f
Year 2	1.33	0.00	f	f
Year 3	1.53**	0.00	f	f
Genitourinary system				
Year 1	1.14	1.03	f	f
Year 2	1.32	2.00	f	f
Year 3	1.98***	6.17**	f	f
Musculoskeletal system and connective tissue (reference group)	. . .	. . .	. . .	. . .
Observations	21,671	6,061	9,888	1,091

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Covers cases with cessation decisions reached in FMRs conducted in 2003 or 2008 (and, for former SSI-only recipients, 2007), and followed through 2010.

. . . = not applicable.

* = statistically significant at the 0.1 level.

** = statistically significant at the 0.05 level.

*** = statistically significant at the 0.01 level.

a. Data for 2007 and 2008 are pooled because of the small SSI-only sample size for 2008.

b. Included as a time-varying effect because the CIF did not satisfy the proportionality assumption. See lower panel.

c. Sample size too small to permit statistically meaningful estimates.

d. Categories were pooled to provide a sample large enough to permit statistically meaningful estimates.

e. Impairment type missing from CDR Waterfall data file.

f. No time-varying Cox regression was calculated because the CIF shown in the upper panel satisfied the proportionality assumption.

hazard ratios are statistically significant at commonly accepted levels and even fewer are significant in the 2008 and 2007/2008 regressions. However, this is likely due to small sample sizes, leaving us unable to determine the extent to which the hazard ratios have changed over time.

Conclusion

In this article, we provide data to address the question: Do individuals who lose disability benefits because of medical improvement return to DI or SSI? We estimate that for adults whose program eligibility ceased because of medical improvement, 30 percent of former SSI-only recipients and 22 percent of former concurrent workers will return to SSI within 8 years. We estimate that about 20 percent of former DI-only workers and about 21 percent of former concurrent workers will return to DI within 8 years of the cessation decision.

Our empirical models use several variables that are also used by SSA in the profiling model that predicts the likelihood of medical recovery and therefore determines who receives a FMR. Thus, the CDR profile score is highly significant in our empirical models for former DI-only workers and, to some extent, for former concurrent workers who return to SSI. In our view, that result demonstrates the usefulness of the profiling model not just for determining who is likely to improve medically at the time of the FMR, but also who is likely to stay off the program in the future.

If funding restricts the number of CDRs to less-than-optimal levels, then some individuals whose eligibility could have ceased will instead continue receiving benefits. Against that scenario, increasing the number of CDRs would likely increase overall savings. However, the program return rate for individuals receiving those additional CDRs could exceed that for individuals undergoing current (restricted-level) CDRs within a particular type of CDR (for example, DI worker, SSI adult, SSI child); in that case, the cost/savings ratio would decline. To understand why, consider that beneficiaries whose eligibility ceases are among the least likely to have a severe disability. Thus, if the number of CDRs within a particular category were to increase above current restricted levels, then beneficiaries losing eligibility in CDRs they otherwise would not receive are likely to have somewhat more severe disabilities, and be somewhat more likely to return to the program, than those losing eligibility in current-level CDRs. Consequently, an increase in certain CDRs could lead to a higher program return rate within that category, thereby decreasing the savings per dollar spent even though overall program savings would still increase. It is important to reiterate that savings per dollar spent is highly dependent on the composition of CDR types as well as assumptions regarding interest rates and cost-of-living adjustments.

By limiting our analysis to post-FMR outcomes before age 62, our results likely describe a lower bound on program return. Individuals may be eligible for SSI based on their disability (and income and resources) until they reach age 65; thereafter, the disability requirement no longer applies. Similarly, individuals can receive DI benefits until they reach their full retirement age. Eligibility at those older ages may be amplified by worsening health. Thus, some individuals in our target population may still return to SSI or DI after what we termed early retirement; however, relatively few people reach age 62 during our observation period, so the effect of those sample restrictions on our estimates may be of little import.

One broader concern not considered in this article is the general health of individuals whose disability program participation ceases because they have medically improved to the point where they no longer meet SSA's eligibility requirements. Such individuals may still have substantial disabilities and limitations. We also cannot tell if those who return to the programs do so because their original disability worsens, or if they reapply because of a new disabling condition.[37]

This article focused on a program-integrity aspect of FMRs. Although most formerly eligible individuals remain off the program, we did not consider their economic situation. Future research should examine the extent to which formerly eligible beneficiaries and recipients reenter the labor force. The availability of employment opportunities likely affects program return. Additionally, further exploration of income (especially at the family level) and use of other programs (for example, vocational rehabilitation) for formerly eligible beneficiaries may also shed more light on why some individuals return to the program and others do not.

Table A-1.
Sample sizes and selection procedures

Restriction	DI-only workers		SSI-only recipients		Concurrent workers	
	Number	Percent	Number	Percent	Number	Percent
CDR Waterfall file extract (2003–2008)	598,728	100.00	571,003	100.00	320,412	100.00
Individuals removed from sample						
CDR profile sample or expedited reinstatement cases[a]	111,234	18.58	96,617	16.92	46,394	14.48
Final FMR decision is missing or precedes initial FMR decision	37	0.01	37	0.01	31	0.01
Died before final FMR decision	1,005	0.17	798	0.14	445	0.14
Awaiting appeal decision or still has time to appeal	682	0.11	163	0.03	479	0.15
Reached aged 60 before initial FMR decision or age 62 before final FMR decision	11,071	1.85	12,703	2.22	2,748	0.86
Eligibility did not cease or CDR profile score is missing	441,323	73.71	436,171	76.39	252,573	78.83
Final sample size	33,376	5.57	24,514	4.29	17,742	5.54

SOURCE: Authors' calculations using Social Security administrative records.

NOTE: Rounded components of percentage distributions do not necessarily sum to 100.

a. Expedited reinstatement cases are actually FMRs for individuals who have had their benefits ceased and are filing for benefits through an expedited process under which they must undergo a FMR to have benefits reinstated.

Table B-1.
Descriptive characteristics of adult DI-only workers whose FMRs resulted in eligibility cessation, 2003 and 2008

Characteristic	2003		2008		Change	
	Number	Percent	Number	Percent	Percent-age points	Percent
Total	7,582	100.00	2,222	100.00	. . .	. . .
Diary type (prospective medical improvement)						
Not expected	264	3.48	139	6.26	2.78	79.89
Possible	5,142	67.82	1,652	74.35	6.53	9.63
Expected	2,176	28.70	431	19.40	-9.30	-32.40
CDR profile score						
Low	553	7.29	173	7.79	0.50	6.86
Medium	950	12.53	601	27.05	14.52	115.88
High	6,079	80.18	1,448	65.17	-15.01	-18.72
Age at initial CDR decision						
Younger than 30	679	8.96	153	6.89	-2.07	-23.10
30–39	2,483	32.75	660	29.70	-3.05	-9.31
40–49	3,216	42.42	971	43.70	1.28	3.02
50–59	1,204	15.88	438	19.71	3.83	24.12
Years in program						
Fewer than 2	87	1.15	(X)	(X)	(X)	(X)
2–3	2,552	33.66	187	8.42	-25.24	-74.99
4–5	2,169	28.61	573	25.79	-2.82	-9.86
6 or more	2,774	36.59	1,462	65.80	29.21	79.83

(Continued)

Table B-1.
Descriptive characteristics of adult DI-only workers whose FMRs resulted in eligibility cessation, 2003 and 2008—*Continued*

Characteristic	2003		2008		Change	
	Number	Percent	Number	Percent	Percent-age points	Percent
Diagnosis						
Neoplasms	913	12.04	208	9.36	-2.68	-22.26
Intellectual disabilities	95	1.25	36	1.62	0.37	29.60
Schizophrenia and other psychotic disorders	401	5.29	168	7.56	2.27	42.91
Other mental disorders	2,117	27.92	725	32.63	4.71	16.87
Diseases of the—						
Endocrine, nutritional, and metabolic system	183	2.41	32	1.44	-0.97	-40.25
Nervous system and sense organs	424	5.59	118	5.31	-0.28	-5.01
Circulatory system	304	4.01	98	4.41	0.40	9.98
Respiratory system	116	1.53	40	1.80	0.27	17.65
Digestive system	360	4.75	85	3.83	-0.92	-19.37
Genitourinary system	407	5.37	132	5.94	0.57	10.61
Musculoskeletal system and connective tissue	1,121	14.79	369	16.61	1.82	12.31
Injuries	820	10.82	130	5.85	-4.97	-45.93
Other	142	1.87	43	1.94	0.07	3.74
Unknown [a]	179	2.36	38	1.71	-0.65	-27.54
Mailer receipt status						
No (direct release to FMR)	6,358	83.86	952	42.84	-41.02	-48.91
Yes	1,224	16.14	1,270	57.16	41.02	254.15
Adjudication level of initial program entry						
Initial application	311	4.10	47	2.12	-1.98	-48.29
Reconsideration	4,815	63.51	1,394	62.74	-0.77	-1.21
Administrative Law Judge or higher	758	10.00	227	10.22	0.22	2.20
Unknown	1,698	22.40	554	24.93	2.53	11.29
Prior CDR status						
No	6,150	81.11	1,626	73.18	-7.93	-9.78
Yes	1,432	18.89	596	26.82	7.93	41.98
Consultative examination request status						
No	4,396	57.98	1,082	48.69	-9.29	-16.02
Yes	3,186	42.02	1,140	51.31	9.29	22.11
Preeligibility earnings quartile						
Lowest	1,899	25.05	607	27.32	2.27	9.06
Second	1,952	25.75	592	26.64	0.89	3.46
Third	1,865	24.60	545	24.53	-0.07	-0.28
Highest	1,866	24.61	478	21.51	-3.10	-12.60

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Rounded components of percentage distributions do not necessarily sum to 100.

. . . = not applicable.

(X) = suppressed to avoid disclosing information about particular individuals.

a. Impairment type is missing in the CDR Waterfall data file.

Table B-2.
Descriptive characteristics of adult SSI-only recipients whose FMRs resulted in eligibility cessation, 2003 and 2008

Characteristic	2003		2008		Change	
	Number	Percent	Number	Percent	Percent-age points	Percent
Total	9,888	100.00	413	100.00	. . .	. . .
Diary type (prospective medical improvement)						
Not expected	433	4.38	71	17.19	12.81	292.47
Possible	7,859	79.48	315	76.27	-3.21	-4.04
Expected	1,596	16.14	27	6.54	-9.60	-59.48
CDR profile score						
Low	661	6.68	47	11.38	4.70	70.36
Medium	1,373	13.89	118	28.57	14.68	105.69
High	7,854	79.43	248	60.05	-19.38	-24.40
Age at initial CDR decision						
Younger than 30	3,671	37.13	85	20.58	-16.55	-80.42
30–39	2,784	28.16	115	27.85	-0.31	-1.11
40–49	2,710	27.41	124	30.02	2.61	8.69
50–59	723	7.31	89	21.55	14.24	66.08
Years in program						
Fewer than 4	1,144	11.57	(X)	(X)	(X)	(X)
4–5	1,988	20.11	(X)	(X)	(X)	(X)
6 or more	6,756	68.33	408	98.79	30.46	30.83
Diagnosis						
Neoplasms	301	3.04	(X)	(X)	(X)	(X)
Intellectual disabilities	1,735	17.55	108	26.15	8.60	32.89
Schizophrenia and other psychotic disorders	873	8.83	57	13.80	4.97	36.01
Other mental disorders	3,330	33.68	159	38.50	4.82	12.52
Diseases of the—						
Endocrine, nutritional, and metabolic system	332	3.36	12	2.91	-0.45	-15.46
Nervous system and sense organs	598	6.05	22	5.33	-0.72	-13.51
Circulatory system	194	1.96	(X)	(X)	(X)	(X)
Respiratory system	216	2.18	(X)	(X)	(X)	(X)
Digestive system	189	1.91	(X)	(X)	(X)	(X)
Genitourinary system	221	2.24	(X)	(X)	(X)	(X)
Musculoskeletal system and connective tissue	557	5.63	13	3.15	-2.48	-78.73
Injuries	333	3.37	(X)	(X)	(X)	(X)
Other	176	1.78	(X)	(X)	(X)	(X)
Unknown [a]	833	8.42	(X)	(X)	(X)	(X)
Mailer receipt status						
No (direct release to FMR)	7,541	76.26	90	21.79	-54.47	-71.43
Yes	2,347	23.74	323	78.21	54.47	229.44
Adjudication level of initial program entry						
Initial application	1,661	16.80	48	11.62	-5.18	-30.83
Reconsideration	6,623	66.98	308	74.58	7.60	11.35
Administrative Law Judge or higher	813	8.22	25	6.05	-2.17	-26.40
Unknown	791	8.00	32	7.75	-0.25	-3.13
Prior CDR status						
No	7,886	79.75	264	63.92	-15.83	-19.85
Yes	2,002	20.25	149	36.08	15.83	78.17

(Continued)

Table B-2.
Descriptive characteristics of adult SSI-only recipients whose FMRs resulted in eligibility cessation, 2003 and 2008—*Continued*

	2003		2008		Change	
Characteristic	Number	Percent	Number	Percent	Percent-age points	Percent
Consultative examination request status						
No	4,700	47.53	228	55.21	7.68	16.16
Yes	5,188	52.47	185	44.79	-7.68	-14.64
Age at initial program entry						
Younger than 18	2,743	27.74	106	25.67	-2.07	-7.46
18 or older	7,145	72.26	307	74.33	2.07	2.86
Preeligibility earnings quartile						
Lowest or second	4,880	49.35	204	49.40	0.05	0.10
Third	2,497	25.25	113	27.36	2.11	8.36
Highest	2,507	25.35	96	23.24	-2.11	-8.32

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Rounded components of percentage distributions do not necessarily sum to 100.

. . . = not applicable.

(X) = suppressed to avoid disclosing information about particular individuals.

a. Impairment type is missing in the CDR Waterfall data file.

Table B-3.
Descriptive characteristics of adult concurrent workers whose FMRs resulted in eligibility cessation, 2003 and 2008

	2003		2008		Change	
Characteristic	Number	Percent	Number	Percent	Percent-age points	Percent
Total	4,022	100.00	1,053	100.00	. . .	. . .
Diary type (prospective medical improvement)						
Not expected	113	2.81	46	4.37	1.56	55.52
Possible	2,667	66.31	761	72.27	5.96	8.99
Expected	1,242	30.88	246	23.36	-7.52	-24.35
CDR profile score						
Low	186	4.62	56	5.32	0.70	15.15
Medium	650	16.16	272	25.83	9.67	59.84
High	3,186	79.21	725	68.85	-10.36	-13.08
Age at initial CDR decision						
Younger than 30	1,029	25.58	232	22.03	-3.55	-16.11
30–39	1,449	36.03	362	34.38	-1.65	-4.80
40–49	1,271	31.60	343	32.57	0.97	2.98
50–59	273	6.79	116	11.02	4.23	38.38
Years in program						
Fewer than 2	123	3.06	10	0.95	-2.11	-222.11
2–3	1,416	35.21	89	8.45	-26.76	-316.69
4–5	1,063	26.43	260	24.69	-1.74	-7.05
6 or more	1,420	35.31	694	65.91	30.60	46.43

(Continued)

Table B-3.
Descriptive characteristics of adult concurrent workers whose FMRs resulted in eligibility cessation, 2003 and 2008—*Continued*

Characteristic	2003		2008		Change	
	Number	Percent	Number	Percent	Percent-age points	Percent
Diagnosis						
Neoplasms	279	6.94	56	5.32	-1.62	-30.45
Intellectual disabilities	220	5.47	86	8.17	2.70	33.05
Schizophrenia and other psychotic disorders	347	8.63	154	14.62	5.99	40.97
Other mental disorders	1,481	36.82	411	39.03	2.21	5.66
Diseases of the—						
Endocrine, nutritional, and metabolic system	78	1.94	21	1.99	0.05	2.51
Nervous system and sense organs	251	6.24	52	4.94	-1.30	-26.32
Circulatory system	132	3.28	27	2.56	-0.72	-28.13
Respiratory system	67	1.67	17	1.61	-0.06	-3.73
Digestive system	119	2.96	21	1.99	-0.97	-48.74
Genitourinary system	124	3.08	30	2.85	-0.23	-8.07
Musculoskeletal system and connective tissue	410	10.19	89	8.45	-1.74	-20.59
Injuries	365	9.08	58	5.51	-3.57	-64.79
Other	80	1.99	21	1.99	0.00	0.00
Unknown [a]	69	1.72	10	0.95	-0.77	-81.05
Mailer receipt status						
No (direct release to FMR)	3,636	90.40	653	62.01	-28.39	-31.40
Yes	386	9.60	400	37.99	28.39	295.73
Adjudication level of initial program entry						
Initial application	197	4.90	31	2.94	-1.96	-40.00
Reconsideration	2,511	62.43	705	66.95	4.52	7.24
Administrative Law Judge or higher	503	12.51	105	9.97	-2.54	-20.30
Unknown	811	20.16	212	20.13	-0.03	-0.15
Prior CDR status						
No	3,164	78.67	737	69.99	-8.68	-11.03
Yes	858	21.33	316	30.01	8.68	40.69
Consultative examination request status						
No	2,111	52.49	496	47.10	-5.39	-10.27
Yes	1,911	47.51	557	52.90	5.39	11.34
Preeligibility earnings quartile						
Lowest	956	23.77	320	30.39	6.62	27.85
Second	1,010	25.11	289	27.45	2.34	9.32
Third	1,058	26.31	254	24.12	-2.19	-8.32
Highest	998	24.81	190	18.04	-6.77	-27.29

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Rounded components of percentage distributions do not necessarily sum to 100.

. . . = not applicable.

a. Impairment type is missing in the CDR Waterfall data file.

Table C-1.
Standard errors for proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Aggregate effects			
Diary type (prospective medical improvement)						
Not expected	0.13	0.17	0.42	0.09	0.17	a
Possible	0.04	0.07	0.12	0.05	0.06	a
Expected (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
CDR profile score						
Low (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Medium	a	a	0.17	0.06	a	0.09
High	a	a	0.24	0.06	a	0.09
Age at initial CDR decision						
Younger than 30 (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
30–39	0.08	0.10	0.07	0.05	a	0.08
40–49	0.11	0.13	0.09	0.08	a	0.12
50–59	0.16	0.16	0.18	0.13	a	0.20
Years in program						
Fewer than 2 (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .	. . .	. . .
2–3	0.08	0.17	. . .	. . .	0.09	0.11
Fewer than 4 (reference group for SSI-only)	. . .	. . .	. . .	. . .	. . .	. . .
4–5	0.09	0.18	0.10	0.06	0.10	0.13
6 or more	0.07	0.17	0.08	0.04	0.10	0.12
Diagnosis						
Neoplasms	0.06	0.08	0.18	0.06	0.09	0.09
Intellectual disabilities	0.17	0.23	0.21	0.09	0.13	0.13
Schizophrenia and other psychotic disorders	0.13	0.18	0.23	0.10	0.14	0.13
Other mental disorders	0.06	0.09	0.16	0.07	0.08	0.08
Diseases of the—						
Endocrine, nutritional, and metabolic system	0.13	0.18	0.37	0.12	0.18	0.15
Nervous system and sense organs	0.08	0.11	0.23	0.09	0.10	0.10
Circulatory system	0.10	0.15	0.32	0.12	0.13	0.13
Respiratory system	0.14	0.19	0.22	0.11	0.17	0.17
Digestive system	0.08	0.11	0.24	0.10	0.11	0.11
Genitourinary system	0.11	0.15	0.47	0.13	0.17	0.15
Musculoskeletal system and connective tissue (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Injuries	0.06	0.09	0.21	0.09	0.07	0.08
Other	0.12	0.18	0.38	0.13	0.18	0.18
Unknown [b]	0.11	0.14	0.22	0.08	0.15	0.15
Mailer receipt status						
No (direct release to FMR; reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes	0.05	a	0.06	0.04	0.07	0.07
Adjudication level of initial program entry						
Initial application (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Reconsideration	0.05	0.06	0.10	0.05	0.07	0.07
Administrative Law Judge or higher	0.04	0.06	0.11	0.05	0.06	0.05
Unknown	0.10	0.14	0.10	0.04	0.12	0.10

(Continued)

Table C-1.
Standard errors for proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Aggregate effects (cont.)			
Prior CDR status						
No (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes	a	a	0.08	0.02	a	a
Consultative examination request status						
No (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes	0.03	0.04	0.06	0.02	0.04	0.03
Age at initial program entry (SSI only)						
Younger than 18 (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
18 or older	. . .	. . .	0.13	0.05	. . .	. . .
Preeligibility earnings quartile						
Lowest (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .	. . .	. . .
Second	a	0.04	. . .	. . .	a	0.06
Lowest or second (reference group for SSI-only)	. . .	. . .	. . .	. . .	. . .	. . .
Third	a	0.03	0.10	0.03	a	0.05
Highest	a	0.02	0.14	0.03	a	0.05
			Time-varying effects			
Diary type (prospective medical improvement)						
Not expected						
Year 1	c	c	c	c	c	0.30
Year 2	c	c	c	c	c	0.30
Year 3	c	c	c	c	c	0.26
Year 4	c	c	c	c	c	0.34
Year 5	c	c	c	c	c	0.56
Year 6	c	c	c	c	c	0.31
Year 7 or 8	c	c	c	c	c	0.27
Possible						
Year 1	c	c	c	c	c	0.11
Year 2	c	c	c	c	c	0.10
Year 3	c	c	c	c	c	0.12
Year 4	c	c	c	c	c	0.13
Year 5	c	c	c	c	c	0.15
Year 6	c	c	c	c	c	0.23
Year 7 or 8	c	c	c	c	c	0.21
Expected (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
CDR profile score						
Low (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Medium						
Year 1	0.10	0.11	c	c	0.39	c
Year 2	0.08	0.13	c	c	0.15	c
Year 3	0.11	0.13	c	c	0.16	c
Year 4	0.22	0.20	c	c	0.70	c
Year 5	0.27	0.41	c	c	0.31	c
Year 6	0.34	0.20	c	c	0.09	c
Year 7 or 8	0.27	0.90	c	c	0.92	c

(Continued)

Table C-1.
Standard errors for proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Time-varying effects (cont.)			
CDR profile score (cont.)						
High						
Year 1	0.08	0.13	c	c	0.27	c
Year 2	0.08	0.12	c	c	0.15	c
Year 3	0.10	0.13	c	c	0.17	c
Year 4	0.18	0.15	c	c	0.77	c
Year 5	0.19	0.48	c	c	0.29	c
Year 6	0.34	0.16	c	c	0.15	c
Year 7 or 8	0.22	0.74	c	c	0.56	c
Age at initial CDR decision						
Younger than 30 (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
30–39						
Year 1	c	c	c	c	c	0.17
Year 2	c	c	c	c	c	0.16
Year 3	c	c	c	c	c	0.15
Year 4	c	c	c	c	c	0.18
Year 5	c	c	c	c	c	0.23
Year 6	c	c	c	c	c	0.29
Year 7 or 8	c	c	c	c	c	0.21
40–49						
Year 1	c	c	c	c	c	0.22
Year 2	c	c	c	c	c	0.22
Year 3	c	c	c	c	c	0.23
Year 4	c	c	c	c	c	0.25
Year 5	c	c	c	c	c	0.34
Year 6	c	c	c	c	c	0.52
Year 7 or 8	c	c	c	c	c	0.24
50–59						
Year 1	c	c	c	c	c	0.42
Year 2	c	c	c	c	c	0.42
Year 3	c	c	c	c	c	0.46
Year 4	c	c	c	c	c	0.65
Year 5	c	c	c	c	c	0.78
Year 6	c	c	c	c	c	1.14
Year 7 or 8	c	c	c	c	c	0.26
Mailer receipt status						
No (direct release to FMR; reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes						
Year 1	c	0.13	c	c	c	c
Year 2	c	0.12	c	c	c	c
Year 3	c	0.15	c	c	c	c
Year 4	c	0.10	c	c	c	c
Year 5	c	0.20	c	c	c	c
Year 6	c	0.20	c	c	c	c
Year 7 or 8	c	0.32	c	c	c	c

(Continued)

Table C-1.
Standard errors for proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 8 years of a 2003–2008 FMR cessation decision, by former program type and beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, return to—		Former SSI-only recipients, return to—		Former concurrent workers, return to—	
	DI	SSI	DI	SSI	DI	SSI
			Time-varying effects (cont.)			
Prior CDR status						
No (reference group)	. . .	. . .	. . .	. . .	. . .	. . .
Yes						
Year 1	0.03	0.03	c	c	0.03	0.03
Year 2	0.04	0.04	c	c	0.04	0.04
Year 3	0.04	0.05	c	c	0.06	0.06
Year 4	0.05	0.07	c	c	0.07	0.06
Year 5	0.05	0.08	c	c	0.09	0.07
Year 6	0.09	0.12	c	c	0.13	0.11
Year 7 or 8	0.08	0.09	c	c	0.14	0.12
Preeligibility earnings quartile						
Lowest (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .	. . .	. . .
Second						
Year 1	0.09	c	. . .	. . .	0.14	c
Year 2	0.10	c	. . .	. . .	0.13	c
Year 3	0.09	c	. . .	. . .	0.20	c
Year 4	0.12	c	. . .	. . .	0.25	c
Year 5	0.15	c	. . .	. . .	0.29	c
Year 6	0.36	c	. . .	. . .	0.42	c
Year 7 or 8	0.25	c	. . .	. . .	0.40	c
Third						
Year 1	0.08	c	c	c	0.14	c
Year 2	0.08	c	c	c	0.12	c
Year 3	0.08	c	c	c	0.23	c
Year 4	0.11	c	c	c	0.26	c
Year 5	0.12	c	c	c	0.32	c
Year 6	0.27	c	c	c	0.37	c
Year 7 or 8	0.31	c	c	c	0.58	c
Highest						
Year 1	0.07	c	c	c	0.15	c
Year 2	0.07	c	c	c	0.12	c
Year 3	0.07	c	c	c	0.19	c
Year 4	0.08	c	c	c	0.33	c
Year 5	0.12	c	c	c	0.28	c
Year 6	0.28	c	c	c	0.30	c
Year 7 or 8	0.18	c	c	c	0.91	c
Observations	169,466	175,582	24,522	24,522	87,854	87,437

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Covers cessation decisions reached in FMRs conducted in 2003–2008, and followed through 2010.

. . . = not applicable.

a. Included as a time-varying effect.

b. Impairment type missing from CDR Waterfall data file.

c. No time-varying Cox regression was calculated because the CIF satisfied the proportionality assumption.

Table D-1.
Standard errors for proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 3 years of a 2003 or 2008 FMR cessation decision, by selected beneficiary characteristics

Characteristic	Former DI-only workers, returned to DI within 3 years of FMR in—		Former SSI-only recipients, returned to SSI within 3 years of FMR in—	
	2003	2008	2003	2007–2008 [a]
	Aggregate effects			
Diary type (prospective medical improvement)				
Not expected	0.18	0.74	0.12	0.52
Possible	0.07	0.30	0.06	0.45
Expected (reference group)	. . .	. . .	. . .	. . .
CDR profile score				
Low (reference group)	. . .	. . .	. . .	. . .
Medium	0.12	0.22	0.10	0.28
High	a	a	0.10	0.28
Age at initial CDR decision				
Younger than 30 (reference group)	. . .	. . .	. . .	. . .
30–39	a	a	0.08	0.20
40–49	0.19	0.89	0.11	0.25
50–59	0.28	1.48	0.18	0.51
Years in program				
Fewer than 4 (reference group for DI-only)	. . .	. . .	. . .	. . .
4–5	0.07	0.38	. . .	. . .
Fewer than 6 (reference group for SSI-only)	. . .	. . .	. . .	. . .
6 or more	0.09	0.29	0.04	0.34
Diagnosis				
Neoplasms	0.11	0.27	b	b
Intellectual disabilities	0.38	1.84	0.08	0.32
Schizophrenia and other psychotic disorders	0.24	0.41	[c] 0.06	[c] 0.23
Other mental disorders	0.11	0.26	[c] 0.06	[c] 0.23
Diseases of the—				
Endocrine, nutritional, and metabolic system				
Nervous system and sense organs	0.14	0.31	0.09	0.40
Circulatory system	0.20	0.64	b	b
Respiratory system	0.24	0.86	b	b
Digestive system	0.12	0.25	b	b
Genitourinary system				
Musculoskeletal system and connective tissue (reference group)	. . .	. . .	. . .	. . .
Injuries	0.09	0.39	b	b
Other	0.20	0.58	b	b
Unknown [d]	0.19	1.08	b	b
Mailer receipt status				
No (direct release to FMR; reference group)	. . .	. . .	. . .	. . .
Yes	0.14	0.20	0.07	0.26

(Continued)

Table D-1.
Standard errors for proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 3 years of a 2003 or 2008 FMR cessation decision, by selected beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, returned to DI within 3 years of FMR in—		Former SSI-only recipients, returned to SSI within 3 years of FMR in—	
	2003	2008	2003	2007–2008 [a]
		Aggregate effects (cont.)		
Adjudication level of initial program entry				
Initial application (reference group)	. . .	. . .	. . .	. . .
Reconsideration	0.09	0.23	0.07	0.26
Administrative Law Judge or higher	0.07	0.27	0.08	0.21
Unknown	0.13	0.78	0.05	0.27
Prior CDR status				
No (reference group)	. . .	. . .	. . .	. . .
Yes	0.04	0.05	0.04	0.11
Consultative examination request status				
No (reference group)	. . .	. . .	. . .	. . .
Yes	0.06	0.17	0.04	0.15
Age at initial program entry (SSI only)				
Younger than 18 (reference group)	. . .	. . .	. . .	. . .
18 or older	. . .	. . .	0.07	0.36
Preeligibility earnings quartile				
Lowest (reference group for DI-only and concurrent)	. . .	. . .	. . .	. . .
Second	0.10	0.25	. . .	. . .
Lowest or second (reference group for SSI-only)	. . .	. . .	. . .	. . .
Third	0.09	0.22	0.05	0.26
Highest	0.07	0.21	0.05	0.24
		Time-varying effects		
CDR profile score				
Low (reference group)	. . .	. . .	. . .	. . .
High				
Year 1	0.14	0.31	e	e
Year 2	0.23	0.20	e	e
Year 3	0.20	0.10	e	e
Age at initial CDR decision				
Younger than 30 (reference group)	. . .	. . .	. . .	. . .
30–39				
Year 1	0.19	0.43	e	e
Year 2	0.21	0.73	e	e
Year 3	0.14	1.92	e	e

(Continued)

Table D-1.
Standard errors for proportional hazard regression results (hazard ratios) of time to first successful reapplication to DI or SSI within 3 years of a 2003 or 2008 FMR cessation decision, by selected beneficiary characteristics—*Continued*

Characteristic	Former DI-only workers, returned to DI within 3 years of FMR in—		Former SSI-only recipients, returned to SSI within 3 years of FMR in—	
	2003	2008	2003	2007–2008 [a]
	Time-varying effects (cont.)			
Diagnosis				
Diseases of the—				
Endocrine, nutritional, and metabolic system				
Year 1	0.20	1.95	b	b
Year 2	0.48	0.00	b	b
Year 3	0.32	0.00	b	b
Genitourinary system				
Year 1	0.33	0.56	b	b
Year 2	0.33	1.25	b	b
Year 3	0.30	4.73	b	b
Musculoskeletal system and connective tissue (reference group)	. . .	. . .	. . .	. . .
Observations	21,671	6,061	9,888	1,091

SOURCE: Authors' calculations using Social Security administrative records.

NOTES: Covers cessation decisions reached in FMRs conducted in 2003 or 2008 (and, for former SSI-only recipients, 2007), and followed through 2010.

. . . = not applicable.

a. SSI data for 2008 are available only in combination with 2007 data.

b. Sample size too small to permit statistically meaningful estimates.

c. Categories were pooled to provide a sample large enough to permit statistically meaningful estimates.

d. Impairment type missing from CDR Waterfall data file.

e. No time-varying Cox regression was calculated because the CIF satisfied the proportionality assumption.

Notes

Acknowledgments: The authors are grateful to the following people for their comments on earlier versions of this article: David Weaver, Richard Balkus, Robert Weathers, Susan Wilschke, Susan Grad, Scott Muller, John Hennessey, Tom Hale, Eli Donkar, Michael Stephens, Ray Wise, Perry Cocke, John Tambornino, Manasi Deshpande, Timothy Moore, MJ Pencarski, Susan Kalasunas, Art Spencer, Judith O'Malley-Oceanak, and especially Bob Somers for his comments and assistance with the data.

[1] The SGA earnings level for 2013 is $1,040. To be eligible for SSI, an individual is limited to $2,000 in countable resources. Once receiving SSI payments, an individual must continue to meet the resource limit but can have earnings above the SGA level. Payments are reduced $1 for every $2 earned above $65 in a month. Unearned income, such as DI benefits, is offset dollar-for-dollar after the first $20. Additional exclusions to income and assets factor into the determination of the monthly SSI payment and optional state supplemental payments. Most SSI recipients are also Medicaid participants. SSI also provides payments to individuals aged 65 or older without disabilities, although the income and asset limits still apply. See SSA (2012f) for more information on SSI rules.

[2] Insured status for DI requires an individual to have a sufficient work history, measured in quarters of coverage, over a recent period. In 2013, an individual earns one quarter of coverage for each $1,160 earned and may earn up to four quarters of coverage per year. For younger workers, fewer quarters of coverage are required to reach insured status. Individuals awarded DI benefits receive a monthly benefit check, as do certain dependent spouses, children, and parents. After 24 months, DI beneficiaries are eligible for Medicare. See SSA (2012c) for more information on DI rules.

[3] The sequential evaluation process used in a CDR, the Medical Improvement Review Standard, differs from that used in an initial disability claim. In general, the review standard process compares the beneficiary's current impairment with that examined at the most recent favorable decision to determine if medical improvement has occurred. Even with evidence of improvement, the examiner must still determine if the severity of the impairment precludes SGA. For exceptions to the Medical Improvement Review Standard, see CFR (1996).

[4] The savings rate is highly dependent on the composition of CDR types (for example, DI worker, SSI adult, SSI child), as well as assumptions regarding interest rates and cost-of-living adjustments.

[5] The president's 2012 budget requested an increase in CDR funding and $938 million for program integrity overall (OMB 2011, 163). SSA expected to spend an estimated $756 million for program integrity in fiscal year 2012 (SSA 2012a).

[6] However, some studies have looked at the related issue of SSI recipients and DI beneficiaries who return to work (for example, Bound 1989; Hennessey and Muller 1995; Schimmel, Stapleton, and Song 2010; Liu and Stapleton 2011; and Schimmel and Stapleton 2011). See also Bound and Burkhauser (1999) for an overview of the research on DI and SSI and Mashaw and Reno (1996) for additional information on DI and SSI policy.

[7] Although those studies and ours examine similar demographic characteristics, we focus on CDR characteristics not available in those studies.

[8] The mailer contains six questions about the individual's health, employment, and medical care use in the last 2 years; for more information, see SSA (2012e). We note that mailer respondents have an inherent incentive to understate their health status. Although that incentive exists throughout the disability determination and review processes, the mailer response does not require supporting medical evidence, which may amplify the incentive. Certain beneficiaries and recipients are not eligible for a mailer. For example, all child SSI recipients, including those undergoing age-18 redeterminations, receive a FMR. SSA does not initiate CDRs for SSI recipients and DI beneficiaries participating in the Ticket to Work Program as long as they are making timely progress toward their employment goals.

[9] Postponed reviews may never take place for some individuals whose characteristics change to the extent that their subsequent profiling model score indicates a lower probability of improvement. Other individuals may leave the programs for other reasons (for example, finding work, reaching full retirement age, or dying).

[10] Over our sample period of 2003 through 2008, about 2.7 percent of mailer cases with a low CDR profile score eventually resulted in a scheduled FMR; however, the availability of resources determined whether those FMRs took place.

[11] The DDS requests a consultative examination when current medical evidence is insufficient to make a decision or if there is conflicting medical information.

[12] Beginning at age 50 (or 45 in certain cases), age is added to the other factors (education, work experience, and residual functional capacity) used in determining an individual's ability to work. Because that change makes the medical improvement standard more difficult to meet, fewer FMRs for older beneficiaries result in cessations.

[13] There are four levels of appeal: reconsideration at the DDS level, the Administrative Law Judge level, an Appeals Council, and federal district court. An individual has 60 days to appeal a cessation decision at each level and 10 days to request continued payments after the initial and reconsideration determinations, although SSA may waive those time limits if there is "good cause." In fiscal year 2008, about 67 percent of adult SSI-only initial cessations

were appealed to the reconsideration level, with 69 percent of those overturned. Additionally, over three-fourths of those with a cessation at the reconsideration level appealed that year; over one-third were successful (SSA 2012b).

[14] SSA field office staff may also initiate FMRs if they have reason to believe medical improvement has occurred. However, SSA's central office initiates the vast majority of reviews, following the process described in this article.

[15] The file is created by the Office of Quality Performance and includes data from various SSA systems including 831/832/833 files, the Supplemental Security Record, the Master Beneficiary Record, and files from the Office of Disability Adjudication and Review. The CDR Waterfall file groups individuals into 10 program-participant categories (such as DI disabled-worker beneficiary, SSI child recipient, and so on), according to their status in July before the fiscal year in which the centrally initiated CDR is scheduled. We only use the SSI adult recipient, DI disabled-worker beneficiary, and disabled-worker concurrent SSI-DI beneficiary groups; other target population restrictions are detailed later. Thus, we include individuals receiving DI benefits only on their own record, not as dependents of other beneficiaries; and adult SSI recipients, meaning they either entered SSI after age 18 or continued in the program after an age-18 redetermination.

[16] We use age at the time of the initial decision for consistency with our other measures. We group individuals into four age groups: younger than 30, 30–39, 40–49, and 50–59.

[17] Expedited reinstatement allows individuals whose benefits terminated because of work to return to DI or SSI through an abbreviated process as long as their medical impairments are the same as, or related to, their original disabling impairments.

[18] Furthermore, the relative scarcity of individuals aged 60–62 would have resulted in imprecise estimates and some multicollinearity issues had we included them.

[19] Our target population includes five individuals who had two FMRs that fit our study criteria. Because the number is relatively small, we do not adjust for any serial correlation that may cause.

[20] Attaining age 62 does not affect SSI eligibility, but we use that cutoff to analyze SSI return for consistency across our analyses. Additionally, attaining age 62 may still affect an individual's behavior because of (a) a family member's receipt of benefits or (b) the difference in the definition of "insured status" between the DI and the Old-Age and Survivors Insurance programs. For example, an individual generally must have worked during the last 10 years to qualify for DI (although there are exceptions for younger workers and people with prior periods of disability); there is no such requirement for the old-age program. Future research might explore the return between age 62 and full retirement age more fully.

[21] We follow Coviello and Boggess (2004) and estimate CIFs using the Stata statistical package. See Hosmer, Lemeshow, and May (2008) for more detail on the Kaplan-Meier survival function.

[22] This estimation strategy is not without its drawbacks. The longest outcomes in our study are based on the earliest cohort in our target population. To the extent that subsequent cohorts are more or less likely to return, die, or reach early retirement age, our estimates could be either too high or too low.

[23] See Hosmer, Lemeshow, and May (2008), Singer and Willett (2003), or Allison (2010) for a fuller discussion of this model.

[24] Hosmer, Lemeshow, and May (2008, 209) note that this model is the same as "specifying an interaction between one of the covariates and the stratification variable."

[25] See also Schoenfeld (1982). Operationally (and equivalently), we test that the log hazard-ratio function is constant over time.

[26] After examining those hazard ratios and formally testing the equality of the ratios for each time-specific effect (that is, the hazard ratio of a high CDR profile score in the first year and the hazard ratio of a high CDR profile score in the second year), we determined that some of the time-varying effects could be combined. For example, as will be shown, the effect of having medical improvement deemed as possible is not significant after the first year; we could thus conceivably combine years 2 through 7+ and improve the efficiency of the empirical model. However, because we estimate multiple empirical models, we keep the yearly effects separate for consistency.

[27] Note that the baseline hazard is not directly estimated by the empirical model, but is recoverable.

[28] This model is described in Hosmer, Lemeshow, and May (2008).

[29] We note that many of the changes over time are likely due to changes in the population size, which drops by at least 70 percent for each target population. One exception is the percentage receiving a mailer, which increased in each group by more than 220 percent from 2003 to 2008. We do not present the distribution of each year's cohort by follow-up status (returned, died, reached age 62, or censored) because such a table does not account for the timing of the event and would likely lead to incorrect interpretations if not viewed carefully. However, such a table is available upon request.

[30] An individual may apply or return to DI on another individual's record (as a child or survivor of another beneficiary). However, more than 98 percent of returns to DI by DI-only and concurrent beneficiaries were on their own record. About 87 percent of former SSI-only recipients entering DI did so on their own record.

[31] It is not immediately clear why former DI-only workers would enter SSI. Most would likely retain their DI-insured status over the observed period. However, some would lose their insured status if they did not return to work. That may explain why entering SSI generally took longer than entering DI (the curve for former DI-only workers in Chart 2 is flatter than that in Chart 1). Possibly, those individuals would have been eligible for concurrent SSI payments but had not applied for SSI. Similarly, some may have been in SSI nonpayment status during the month they were selected for a CDR, and thus were categorized as DI-only on a technicality. (That circumstance may also apply to SSI-only individuals, although suspensions of payments are much less common in DI than in SSI.) Alternatively, many individuals may have spent down their assets while receiving DI benefits (or while dealing with the loss of DI), making them newly eligible for SSI.

[32] In similar (unreported) analyses, we estimate that one-third (33.2 percent) of our DI-only group reapply for DI and over one-half (54.2 percent) of the SSI-only group reapply for SSI after 8 years. Dividing those reapplication rates by the return rates we estimated, the respective postparticipation award rates are 59 percent and 55 percent for the DI-only and SSI-only individuals. Those are substantially higher than the initial award rates reported in SSA publications, which range from 31 percent to 36 percent for DI and from 40 percent to 47 percent for SSI over the observation period (SSA 2012c, 2012f).

[33] The small yearly sample sizes preclude estimating CIFs for each characteristic.

[34] The 95-percent confidence intervals overlap for all neighboring cohorts (for example, 2004 and 2005) during their common followup periods.

[35] Those amounts are adjusted to 2009 dollars using the Consumer Price Index-All Urban Consumers. The nominal values are $551 million in 2003 and $307 million in 2008.

[36] Note that the Great Recession would be expected to shift the curves in the opposite direction—with fewer jobs available, we would expect greater return by the later cohorts early in the followup period; we do not observe that result.

[37] SSA systems record no more than two diagnosis codes for an individual; differences between the precessation-decision and the program-return diagnosis codes would not necessarily identify truly new disabilities, especially in cases of high comorbidity. Similarly, worsening health due to the original disability may not be captured in the data if new impairments occur that more readily meet SSA's definition of disability.

References

Allison, Paul D. 2010. *Survival Analysis Using SAS: A Practical Guide, Second Edition.* Cary, NC: SAS Institute, Inc.

Autor, David H., and Mark G. Duggan. 2003. "The Rise in the Disability Rolls and the Decline in Unemployment." *Quarterly Journal of Economics* 118(1): 157–206.

———. 2006. "The Growth in the Social Security Disability Rolls: A Fiscal Crisis Unfolding." *Journal of Economic Perspectives* 20(3): 71–96.

Black, Dan, Kermit Daniel, and Seth Sanders. 2002. "The Impact of Economic Conditions on Participation in Disability Programs: Evidence from the Coal Boom and Bust." *American Economic Review* 92(1): 27–50.

Bound, John. 1989. "The Health and Earnings of Rejected Disability Insurance Applicants." *The American Economic Review* 79(3): 482–503.

Bound, John, and Richard V. Burkhauser. 1999. "Economic Analysis of Transfer Programs Targeted on People with Disabilities." In *Handbook of Labor Economics*, *Vol. 3C,* edited by Orley C. Ashenfelter and David Card, 3417–3528. Amsterdam: Elsevier.

Burkhauser, Richard V., and Mary C. Daly. 2011. *The Declining Work and Welfare of People with Disabilities: What Went Wrong and a Strategy for Change.* Washington, DC: AEI Press.

[CFR] Code of Federal Regulations. 1996. *How We Will Determine Whether Your Disability Continues or Ends.* Title 20—Employees' Benefits, Chapter III—Social Security Administration, Part 404—Federal Old-Age, Survivors and Disability Insurance (1950–), Subpart P—Determining Disability and Blindness, Section 1594. 20 CFR § 404.1594. http://www.gpo.gov/fdsys/granule/CFR-1996-title20-vol2/CFR-1996-title20-vol2-sec404-1594/content-detail.html.

Coviello, Vincenzo, and Mary Boggess. 2004. "Cumulative Incidence Estimation in the Presence of Competing Risks." *Stata Journal* 4(2): 103–112.

Dykacz, Janice M. 1998. "Return of Disabled-Worker Beneficiaries to the DI Program: Some Insights from the New Beneficiary Followup." *Social Security Bulletin* 61(2): 3–11.

Dykacz, Janice M., and John C. Hennessey. 1989. "Postrecovery Experience of Disabled-Worker Beneficiaries." *Social Security Bulletin* 52(9): 42–66.

Gooley, Ted A., Wendy Leisenring, John Crowley, and Barry E. Storer. 1999. "Estimation of Failure Probabilities in the Presence of Competing Risks: New Representations of Old Estimators." *Statistics in Medicine* 18(6): 695–706.

Grambsch, Patricia M., and Terry M. Therneau. 1994. "Proportional Hazards Tests in Diagnostics Based on Weighted Residuals." *Biometrika* 81(3): 515–526.

Hennessey, John C., and Janice M. Dykacz. 1993. "A Comparison of the Recovery Termination Rates of Disabled-Worker Beneficiaries Entitled in 1972 and 1985." *Social Security Bulletin* 56(2): 58–69.

Hennessey, John C., and L. Scott Muller. 1995. "The Effect of Vocational Rehabilitation and Work Incentives on Helping the Disabled-Worker Beneficiary Back to Work." *Social Security Bulletin* 58(1): 15–28.

Hosmer, David W., Stanley Lemeshow, and Susanne May. 2008. *Applied Survival Analysis: Regression Modeling of Time-to-Event Data, Second Edition.* Hoboken, NJ: John Wiley & Sons, Inc.

Kubik, Jeffrey D. 1999. "Incentives for the Identification and Treatment of Children with Disabilities: The Supplemental Security Income Program." *Journal of Public Economics* 73(2): 187–215.

———. 2003. "Fiscal Federalism and Welfare Policy: The Role of States in the Growth of Child SSI." *National Tax Journal* 56(1): 61–79.

Liu, Su, and David C. Stapleton. 2011. "Longitudinal Statistics on Work Activity and Use of Employment Supports for New Social Security Disability Insurance Beneficiaries." *Social Security Bulletin* 71(3): 35–59.

Marubini, Ettore, and Maria Grazia Valsecchi. 1995. *Analysing Survival Data from Clinical Trials and Observational Studies.* Chichester: John Wiley & Sons Ltd. (UK).

Mashaw, Jerry L., and Virginia P. Reno, editors. 1996. *The Environment of Disability Income Policy: Programs, People, History and Context.* Washington, DC: National Academy of Social Insurance.

[OMB] Office of Management and Budget. 2011. *Fiscal Year 2012 Budget of the U.S. Government.* Washington, DC: Government Printing Office. http://www.gpo.gov/fdsys/pkg/BUDGET-2012-BUD/pdf/BUDGET-2012-BUD.pdf.

Rupp, Kalman, and David C. Stapleton, editors. 1998. *Growth in Disability Benefits: Explanations and Policy Implications.* Kalamazoo, MI: W.E. Upjohn Institute for Employment Research.

Schimmel, Jody, and David C. Stapleton. 2011. "Disability Benefits Suspended or Terminated Because of Work." *Social Security Bulletin* 71(3): 83–103.

Schimmel, Jody, David C. Stapleton, and Jae Song. 2010. "How Common is 'Parking' Among Social Security Disability Insurance (SSDI) Beneficiaries? Evidence from the 1999 Change in the Level of Substantial Gainful Activity (SGA)." MRRC Working Paper No. 2009-220. Ann Arbor, MI: Michigan Retirement Research Center.

Schmidt, Lucie, and Purvi Sevak. 2004. "AFDC, SSI, and Welfare Reform Agressiveness: Caseload Reductions versus Caseload Shifting." *Journal of Human Resources* 39(3): 792–812.

Schmulowitz, Jack. 1973. "Recovery and Benefit Termination: Program Experience of Disabled-Worker Beneficiaries." *Social Security Bulletin* 36(6): 3–15.

Schoenfeld, David. 1982. "Partial Residuals for the Proportional Hazards Regression Model." *Biometrika* 69(1): 239–241.

Singer, Judith D., and John B. Willett. 2003. *Applied Longitudinal Data Analysis: Modeling Change and Event Occurrence.* Oxford: Oxford University Press (UK).

[SSA] Social Security Administration. 2012a. *Annual Performance Plan for FY 2013 and Revised Final Performance Plan for FY 2012.* Washington, DC: SSA. http://www.socialsecurity.gov/performance/2013/.

———. 2012b. *Annual Report of Continuing Disability Reviews, Social Security Administration, Fiscal Year 2010.* http://www.socialsecurity.gov/legislation/FY%202010%20CDR%20Report.pdf.

———. 2012c. *Annual Statistical Report on the Social Security Disability Insurance Program, 2011.* Washington, DC: SSA. http://www.socialsecurity.gov/policy/docs/statcomps/di_asr/2011/index.html.

———. 2012d. *Annual Statistical Supplement to the Social Security Bulletin, 2011.* Washington, DC: SSA. http://www.socialsecurity.gov/policy/docs/statcomps/supplement/2011/index.html.

———. 2012e. "Program Operations Manual System (POMS) Section DI 28001.003: An Overview of Processing Continuing Disability Review (CDR) Mailer Forms SSA-455 and SSA-445-OCR-SM." https://secure.ssa.gov/apps10/poms.nsf/lnx/0428001003.

———. 2012f. *SSI Annual Statistical Report, 2011.* Washington, DC: SSA. http://www.socialsecurity.gov/policy/docs/statcomps/ssi_asr/2011/index.html.

Treitel, Ralph. 1979. "Recovery of Disabled Beneficiaries: A 1975 Followup Study of 1972 Allowances." *Social Security Bulletin* 42(4): 3–23.

Wamhoff, Steve, and Michael Wiseman. 2005/2006. "The TANF/SSI Connection." *Social Security Bulletin* 66(4): 21–36.

Outcome Variation in the Social Security Disability Insurance Program: The Role of Primary Diagnoses

by Javier Meseguer*

Based on the adjudicative process, the author classifies claimant-level data over an 8-year period (1997–2004) into four mutually exclusive categories: (1) initial allowances, (2) initial denials not appealed, (3) final allowances, and (4) final denials. The ability to predict those outcomes is explored within a multilevel modeling framework, with applicants clustered by state and primary diagnosis code. Variance decomposition suggests that medical diagnoses play a substantial role in explaining individual-level variation in initial allowances. Moreover, there is statistically significant high positive correlation between the predictions of an initial allowance and a final allowance across the diagnoses. This finding suggests that the ordinal ranking of impairments between these two adjudicative outcomes is widely preserved. In other words, impairments with a higher expectation of an initial allowance also tend to have a higher expectation of a final allowance.

Introduction

The purpose of the Disability Insurance (DI) program is to replace part of a worker's earnings in the eventuality of a physical or mental impairment preventing the individual from working. The disability portion of the Old-Age, Survivors, and Disability Insurance (OASDI) program, administered by the Social Security Administration (SSA), protects workers and their eligible dependents against such risk. SSA administers a second program, Supplemental Security Income (SSI), which has no employment or contribution requirements, but imposes strict income and asset limits. It is designed to be a program of last resort, assisting aged, blind, or disabled individuals who have very limited resources.

The goal of this study is to explore the extent to which medical diagnoses and state of origin may explain observed heterogeneity in disability decisions. One instance of heterogeneity is manifest at the state level. The DI program is federally administered and is operated in collaboration with the states. When a local Social Security field office establishes that an applicant meets all of his or her nonmedical requirements, the case is forwarded to the state Disability Determination Service (DDS) for a decision. The DDS follows a sequential process to evaluate the medical evidence and decide if the applicant meets the definition of disability. In doing so, a DDS examiner considers the severity of the impairment(s), along with vocational factors that take into account age, education, and work experience. SSA guidelines to determine disability are uniform across all 50 states. In practice, however, there can be wide variation in state allowance rates.

A second instance of variation in DI outcomes occurs through the adjudicative process. If a disability claim is denied, the applicant has a number of opportunities to appeal the decision. There are three stages of appeal within SSA: (1) a reconsideration by the

Selected Abbreviations

ALJ	administrative law judge
DDS	Disability Determination Service
DI	Disability Insurance
DIC	deviance information criterion
DRF	Disability Research File
RFC	residual functional capacity

* *Javier Meseguer is an economist with the Office of Economic Analysis and Comparative Studies, Office of Research, Evaluation, and Statistics, Office of Retirement and Disability Policy, Social Security Administration.*

Selected Abbreviations—*Continued*

SGA	substantial gainful activity
SSA	Social Security Administration
SSI	Supplemental Security Income

state DDS, (2) a hearing before an administrative law judge (ALJ), and (3) a review by the Appeals Council. If those stages are exhausted, the claimant can always seek legal redress in a federal district court. While few initial denials are reversed at the reconsideration level, a substantial portion of claimants who appeal at the hearing level or above are eventually allowed.

The two referenced sources of variation in disability outcomes (by state and adjudicative level) have been a cause of concern to SSA and Congress regarding the practical implementation of the disability programs. My hunch is that the collection of impairments in particular might shed some light in explaining a portion of the observed variation. Thus, I investigate heterogeneity in disability outcomes along three dimensions: state of origin, medical diagnosis, and adjudicative stage. That objective is pursued by working with a random sample of the Disability Research File (DRF). The DRF follows a cohort of applicants through the various stages of the determination process, identifying decisions made at different adjudicative levels. The disability determinations in the file are separated into four mutually exclusive categories: (1) initial allowances, (2) initial denials not appealed, (3) final allowances, and (4) final denials. This classification of the data implicitly reduces the adjudicative process to two stages (*initial* and *final*).

The data is fitted to various Bayesian hierarchical multinomial logit specifications, with two different groups or clusters nesting the claimant-level observations. One group is the 50 states. The other group comprises 181 medical impairments, which represent the unique administrative four-digit primary diagnosis codes. This modeling approach offers several advantages. First, the framework is multivariate, meaning that instead of estimating a separate model for each stage, the adjudicative outcomes are estimated jointly. Second, the multilevel or hierarchical nature of the models enables the distinction to be made between claimant-level effects on one hand and state or diagnosis-level effects on the other hand. In other words, I can decompose heterogeneity in the adjudicative outcomes by source into "between-group" and "within-group" variance. For instance, at one end of the spectrum, it is possible that claimants within a state are rather uniform in their characteristics, so that most of the variance in initial allowances is due to unique differences between the states. Alternatively, a large portion of the total variance could be attributed to claimant-level heterogeneity within the states (that is, the states are not that different from one another, but the population within a given state varies greatly in its characteristics). Finally, a third advantage in this modeling approach is the ability to estimate correlation patterns that may exist between the disability adjudicative outcomes.

The next section in this article provides background information about the Social Security disability programs, including the disability determination and appeals processes. I then briefly review some of the literature regarding the modeling of allowance rates. The data and modeling approach are discussed next, emphasizing the observed variation in adjudicative outcomes by such factors as age, diagnosis group, state of origin, and mortality. The inferential results are presented in the following section, where the "goodness-of-fit" of the various models and the "average effect" of various explanatory variables are evaluated and discussed. Two other important issues addressed in this section involve variance decomposition and correlation, where I describe the interpretation and implications of my estimates. The last section concludes with a summary of the main findings.

Social Security Disability Programs

SSA operates two different programs that offer cash benefits to the disabled: the Disability Insurance program, which was enacted in 1956, and the Supplemental Security Income program, which began in 1972. The two programs share the same disability determination process, but have different objectives. DI is funded through payroll tax contributions and is designed to protect workers contributing to the program from earnings losses that are due to impairment. SSI, on the other hand, is not contributory. General revenues fund it, and the main goal of the program is to guarantee a minimal level of income to the poorest of the aged, blind, or disabled population.

The DI program provides benefits to disabled workers who are younger than their respective full retirement ages and to their spouses, surviving disabled spouses, and disabled children, although workers account for the largest share of beneficiaries (typically, over 80 percent of the DI rolls). At the end of 2010, about 8.8 million workers and their dependents were receiving DI benefits and 4.7 million individuals were receiving SSI payments. Under both programs, the definition of disability is one of long-term work disability. It involves the inability to engage in substantial gainful activity (SGA) because of a medically determinable physical or mental impairment that is expected to last at least 12 months or result in death.

Eligibility for DI benefits requires a worker to be insured, younger than his or her full retirement age, and to meet the definition of disability. The applicant must have worked long enough in employment covered by Social Security (approximately 10 years) and recently enough (about 5 of the past 10 years). Those requirements are relaxed for younger applicants who have shorter employment histories. An applicant who is employed must also have monthly earnings below the SGA threshold ($1,640 for a blind person and $1,000 for a nonblind individual in 2010). However, there are no restrictions on nonwage income. Upon approval, benefits are received after a 5-month waiting period from the onset of disability. In addition, the beneficiary is entitled to Medicare coverage after receiving benefits for 2 years.

Disability benefits continue for as long as the beneficiary remains disabled or reaches full retirement age, in which case there is a conversion to retirement benefits. Upon death of a worker, some dependent benefits may convert into survivor benefits. SSA conducts periodic continuing disability reviews (CDRs) to determine if an individual remains disabled. Review frequency depends on the severity and likelihood of improvement of the disability and can range from 6 months from the initial finding to as long as 7 years. A finding that a beneficiary is engaging in SGA will result in termination.[1]

From 1970 through 2009, the number of beneficiaries in the DI program more than tripled, while DI expenditures increased by almost seven times in inflation-adjusted figures (Congressional Budget Office 2010). According to the Social Security Advisory Board (2012a), that expansion can be traced to several factors in addition to an increase in the general population. One factor has been an increase in the share of lower mortality impairments with earlier onset (such as musculoskeletal and mental disorders). Applicants with those types of impairments tend to enter the program at younger ages and remain as beneficiaries for longer periods of time. Another factor has been an increase in female labor force participation. The rapid pace at which women have joined the ranks among workers has considerably expanded the pool of applicants. Indeed, the gender composition of beneficiaries today is much closer to that of the population at large. A third factor has been an increase in earnings replacement rates. Rising income inequality coupled with the average wage indexing of benefits has increased the portion of potential earnings replaced by DI benefits. Younger low-skilled workers in particular have experienced the highest increase in the value of DI benefits at a time of reduced demand for their labor. Exacerbating the gap between potential earnings and disability benefits is a reduction in private health insurance coverage. Eventual access to Medicare after 2 years on the DI rolls may provide an additional enticement to apply.

The Sequential Disability Determination Process

A claimant typically files an application for DI or SSI in a Social Security field office. The field office gathers a variety of information from the applicant regarding entitlement status, impairment(s), and medical records. The disability determination follows a five-step sequential evaluation process that considers employment, medical, and vocational factors, in that order.

- Step 1: If the applicant is employed and earning more than the SGA amount, an SSA employee denies the claim. Otherwise, the field office sends the claim to the DDS.
- Step 2: If a medical impairment (or combination of impairments) is not severe enough to interfere with basic work-related activities for at least 1 year, a DDS examiner denies the claim. Otherwise, the evaluation proceeds to the next step.
- Step 3: Impairments that meet the criteria in SSA's medical listings or are found to be of equal severity result in an allowance determination. Otherwise, the claim is referred to the next step.
- Step 4: An applicant found with the capacity to engage in relevant employment performed in the past is denied. If not, the evaluation proceeds to the next step.
- Step 5: Based on the applicant's residual functional capacity (RFC), age, education, and work experience, the DDS determines if the applicant could engage in other types of employment. If so, the claim is denied. Otherwise, the claim results in a disability finding.

Motivating the sequential disability determination process is a screening strategy designed to deal first with cases that can be easily decided on the basis of fairly objective medical tests. If the claimant does not meet or equal the severity requirements in the listings of impairments, the vocational grid is used to determine whether he or she is disabled. The grid incorporates a combination of the following factors: age, RFC, education, and the skill level involved in past work as well as the degree to which those skills can be transferred to another job. Age is divided along four thresholds (younger than age 50, aged 50–54, aged 55–59, and aged 60 or older). RFC is graded into five different categories that assess the exertional limitations of the filer for work-related activities (sedentary, light, medium, heavy, and very heavy work). For the purpose of the vocational grid, SSA divides educational level into four categories (illiterate or unable to communicate in English, limited education or less, high school graduate or more, and recent education that trained the applicant for a skilled job). Assessment of previous relevant work experience leads to the categories of unskilled, semiskilled, and skilled. Finally, the determination process takes into account whether the skills the applicant learned from a past job can be transferred to a new, similar position.

Lahiri, Vaughan, and Wixon (1995) and Hu and others (2001) used household survey data matched to Social Security's administrative records to model the sequential disability determination process. Their findings indicate that the predictive ability of particular variables is linked to their relevance within the stage of determination. For instance, information on activity limitations and medical variables are significant to steps 2 and 3, while the explanatory power of age, past work, and education are manifest in steps 4 and 5.

The Appeals Process

Within 60 days from the notice of denial, the applicant has a number of sequential chances to appeal the decision. There are four stages of appeal. The first stage is a reconsideration by the state DDS, where the case is reviewed by a different examiner and the applicant has the opportunity to submit additional evidence. The second stage involves the Office of Disability Adjudication and Review (ODAR), where the claimant can request a hearing before an ALJ.[2] The ALJ considers any documentary evidence introduced, evaluates the testimony of the applicant, and witnesses that testimony under oath. The third stage in the appeals process is to request a review by the Appeals Council, which is comprised of a panel of ALJs. The Council may choose to grant, deny, or dismiss the request. Upon review, the Council can uphold, reverse, or modify the decision. It can also send the case back to the ALJ for a new hearing. Finally, if the applicant is dissatisfied with the outcome, the fourth stage available is to appeal the case outside of SSA in a federal district court.

Table 1 presents allowance, denial, and appeal rates for disability determinations made at various adjudicative stages by year of application. The table reflects 100 percent of the determinations for workers applying to the DI program only, excluding concurrent applicants to DI and SSI. Results are shown for the combined 8-year period spanning the random data sample in my modeling effort (applications

Table 1.
Allowance, denial, and appeal counts and rates for disability determinations at various adjudicative levels, by selected years 1997, 2004, and the 1997–2004 period

Count and rate of disability determination	1997	2004	1997–2004
		Initial level	
Number			
Determinations	551,909	736,987	5,151,351
Allowances	228,793	329,523	2,319,171
Denials	323,116	407,464	2,832,180
Appeals	206,148	248,232	1,778,805
Percent			
Allowance rate	41.45	44.71	45.02
Denial rate	58.55	55.29	54.98
Appeal rate	63.80	60.92	62.81
		Reconsideration level	
Number			
Determinations	206,148	248,232	1,778,805
Allowances	33,373	28,707	255,201
Denials	172,775	219,525	1,523,604
Appeals	141,021	185,672	1,288,257
Percent			
Allowance rate	16.19	11.56	14.35
Denial rate	83.81	88.44	85.65
Appeal rate	81.62	84.58	84.55
		Hearing level or above	
Number			
Determinations	141,021	185,672	1,288,257
Allowances	107,539	151,122	1,009,799
Denials	33,482	34,550	278,458
Percent			
Allowance rate	76.26	81.39	78.38
Denial rate	23.74	18.61	21.62

SOURCE: Author's tabulations based on the *Annual Statistical Report on the Social Security Disability Insurance Program, 2008.*

from 1997 through 2004), as well as separately for 2 individual years (the first (1997) and last (2004)).[3] The initial disability allowance rate within the 8-year period considered stands at about 45 percent. Roughly, 63 percent of initial denials are appealed at the reconsideration stage, which results in a fairly small portion of reversals by the DDS (about 14 percent). However, 85 percent of the reconsideration denials are appealed. Once the third and fourth stages in the appeals process are reached (at the hearing level or in a federal court), denials are reversed at a rate of 78 percent. As a result, after the appeals process takes its course, the 45 percent initial disability allowance rate increases to an overall allowance rate of 70 percent.

Multiple factors can contribute to the high reversal rate of initial denials. The most obvious explanation is that many impairments can worsen over time, particularly disorders that are of a degenerative nature. One feature of the DI program is that at every stage of the appeals process the claimant has an opportunity to introduce additional medical evidence. Therefore, it is possible that ALJs are making decisions based on a more extensive information set that was simply not available to state DDS examiners. Moreover, unlike with the DDS appeals procedure, applicants at the hearing level or above are much more likely to retain legal counsel. Claimant representation benefits from detailed knowledge of the rules and process. This can be helpful in developing medical evidence that may include additional symptoms and impairments not claimed at the DDS level. In this context, the Social Security Advisory Board (2001) has made a number of recommendations addressing some of the procedural differences between the adjudicative levels (such as the fact that most claimants lack any face-to-face interaction with an adjudicator until they get to an ALJ hearing). Finally, by its very nature, the appeals process could be inducing a selection bias effect, where only the applicants with the strongest evidence appeal a denial. In fact, one possible route to selection bias is the use of legal counsel. After all, attorneys are likely to prescreen potential clients in order to represent those with the highest probability of an allowance.[4]

Previous Literature

SSA's statutory definition of disability in terms of "ability to work" is inevitably open to subjective judgment on the part of decision makers. In a minority of cases, proof of a specific impairment will qualify the filer for expedited case processing under the Compassionate Allowance (CAL) initiative, based on minimal, but sufficient objective medical information. Roughly, about a third of allowances are decided on the medical evidence alone (step 3), but even physicians may disagree over the interpretation of diagnostic tests. Most claimants are unlikely to neatly fit precisely defined eligibility criteria, and program guidelines can be subject to interpretation. In some instances, federal courts have issued decisions that at least for a while resulted in different disability policies for different parts of the country.[5] Moreover, individuals vary in their ability to withstand pain and in their response to treatment, so that one person facing a specific set of limitations may be able to work, while another may not. Once vocational considerations such as RFC, relevant past work experience, and transferable skills are criteria in the determination process, the decision becomes increasingly complex. For these reasons alone, one would expect some degree of heterogeneity in disability outcomes.

The literature evaluating factors that affect allowance rates in Social Security's disability programs is extremely sparse. More effort has been devoted to investigating the determinants of application rates. Rupp and Stapleton (1995) summarized earlier contributions, while Rupp (2012) discussed more recent work. A growing body of evidence using different methodology and various sources of data suggests that application rates increase with labor market shocks. Higher unemployment reduces the opportunity cost of applying for marginally qualifying individuals, who must weigh their current earnings and future labor opportunities against the present value of benefits. Thus, application rates are expected to rise in response to a labor market shock. Additionally, the increase in marginally qualified applicants is anticipated to produce a decline in allowance rates, as those filers have a harder time qualifying through the determination process.

For over a decade, the Social Security Advisory Board (2001, 2006, 2012a) has been tracking the two main sources of variation in allowance rates referenced in this article (by state and adjudicative stage), calling for a major overhaul to the disability programs. Among its suggestions, the Board advocates strengthening the federal/state arrangement to decrease the large disparities that exist between different states regarding staff salaries, educational requirements, training, and attrition rates. The Board also recommends reforming the hearing process by establishing uniform procedures for claimant representatives;

having the government represented at the ALJ hearing level or above; and closing the record after the ALJ decision, so that cases do not change substantially at each level of appeal.

Using a combination of aggregate time-series and cross-sectional methodology, Rupp and Stapleton (1995) found a positive relationship between the state unemployment rate and both initial applications and awards. Their modeling of allowance rates suggested the presence of lagged effects. Specifically, the authors estimated that a 1 percentage point increase in the unemployment rate was associated with a 1 percent decline in the initial allowance rate in the first and second years following the year in which the unemployment rate changed.

State allowance rates depend on the economic, demographic, and health characteristics of the applicants, which vary among the states. For instance, states with older populations are anticipated to have higher disability allowance rates on average. Older applicants are more likely to qualify because of the higher prevalence of age-related disabilities and the fact that they face less stringent program standards than do younger individuals. Using state-level data over a 3-year period (1997–1999), Strand (2002) estimated that as much as half of the variation in initial allowance rates may have been attributable to state differences in economic and demographic factors. The author found a negative association between filing rates and allowance rates and a statistically significant negative impact of unemployment on allowance rates. Institutional considerations can also play a role in explaining observed heterogeneity in disability outcomes. For instance, Coe and others (2011) found that states with mandated health insurance and longer duration for Unemployment Insurance benefits were associated with lower application rates.

In a recent article, Rupp (2012) used individual-level data over the 1993–2008 period to investigate three factors affecting initial allowance rates: (1) the demographic characteristics of applicants, (2) the diagnostic mix of applicants, and (3) local labor market conditions. The modeling approach involved a binary logit process with fixed-effects for state of origin and year of determination. Explanatory variables included the state unemployment rate and indicators for sex, age group, impairment type,[6] and the presence of a secondary diagnosis code in the data. The author found these three sets of variables statistically significant. All else equal, male and older adult applicants had a higher likelihood of an initial allowance. Likewise, an increase in the state unemployment rate was associated with a decline in the probability of an initial allowance, with the size of the effect changing substantially by body system. The size of the state fixed-effects suggested that a substantial portion of the variation in state initial allowance rates could be attributed to permanent differences among the states.

Keiser (2010) explored the variation in self-reported (as opposed to actual) allowance rates among DDS examiners in three undisclosed states. The study approached the subject of outcome variation in disability decision making from the perspective of the theory of bounded rationality. The surveys mailed to DDS examiners considered a number of factors, including: (1) ideological identification; (2) adherence to conflicting goals (aiding disabled individuals, while protecting US tax payers from fraud); (3) perception about applicants' honesty in representing their limitations; and (4) the expectations of examiners' immediate supervisors (a focus on allowances, denials, or both equally). The model was able to account for only 12 percent of the variation in self-reported allowance rates. One aspect of the study relevant to the objectives here relates to the evidence of a possible policy feedback mechanism. In particular, knowledge of the extent to which ALJs reverse initial denials was found to be a factor in explaining higher reported allowance rates among examiners.

Data and Methodology

The Disability Research File (DRF) is a data file designed to longitudinally track a cohort of filers through 10 years of the disability decision and appeal process. Prompted by concern from Congress regarding the size of the disability rolls, the file—originally built in 1993—is updated once a year, with the 3 most recent years of claims data completely built from scratch. Because of differences in the structure of DI and SSI records (Title II and Title XVI, respectively, under the Social Security Act), two separate files are compiled that draw from multiple administrative data sources in a process that usually takes several months to complete. The file is unique in its ability to provide information about the status of a claim in its progression throughout the adjudicative stages, as well as activity about claimants who file multiple disability applications.

For this study, I work with a 10 percent random sample of an abbreviated version of the DRF, tracking 10 years of longitudinal disability claims (1997–2006). The analysis is restricted to medical determinations involving workers aged 18–65 who applied to the DI program during the 8-year period from 1997 through 2004. The latter is the most recent year in the file for which the percentage of pending applications is negligible. Moreover, the focus is on DI medical claims only. In particular, technical denials are excluded because they generally lack the evaluation of any medical evidence.[7] Concurrent applicants to the DI and SSI programs are also excluded, as they represent a unique population that has enough work experience to qualify under DI, but that is poor enough to meet SSI's criteria. A look at the *Annual Statistical Report on the Social Security Disability Insurance Program* (SSA 2009, Tables 60 and 62) validates this decision. Nonconcurrent DI workers systematically experience higher allowance rates at the initial and hearing levels than concurrent workers. Furthermore, Rupp (2012, Table 1) illustrates how the age structure and diagnostic mix of both populations can differ substantially. Concurrent filers tend to be younger and have a much larger share of mental diagnoses. Thus, it seems appropriate to treat DI-only, concurrent, and SSI-only claimants as separate populations.

Formally, the adjudicative-level process can be thought of as a sequential interaction between two parties (Social Security and the applicant). Conditional on a claimant applying to the disability program, Social Security makes a decision to allow or deny. Likewise, conditional on a denial, the applicant decides whether or not to appeal. The sequence continues, with the process ending upon an allowance, a decision not to appeal, or exhaustion of all appeals opportunities. While the appeals decision is always made by the same individual (the applicant), the decision to allow or deny can be made by a field office representative, an examiner at the DDS, an ALJ, or even a federal judge. Complicating matters further is the Prototype program, which breaks the order of the sequence by allowing several states to skip the reconsideration adjudicative level.

This article focuses on the prediction of outcomes as a purely statistical classification problem. I do not model the sequential structure of the decision-making process. For purposes of this study, the disability determinations in the file are separated into four mutually exclusive categories: (1) initial allowances, (2) initial denials not appealed, (3) final allowances, and (4) final denials. This classification of the data implicitly reduces the adjudicative process to two stages. Specifically, the first two categories (initial allowances and initial denials not appealed) represent outcomes at the initial DDS level. The last two categories (final allowances and final denials) result once the applicant decides to stop appealing or exhausts the appeals process. This can occur at the reconsideration DDS level, at the hearing level, or in a federal court. In other words, what triggers the difference between the two adjudicative stages is a decision to appeal an initial denial. However, because of the low allowance rate and high appeal rate at the reconsideration stage (see Table 1), the large majority of decisions falling into the final allowance and final denial categories occur at the hearing level or above.

Table 2 breaks down the count and proportion of sample observations by adjudicative disability outcome. In the top panel of the table, out of a random sample comprising 462,578 observations, 46.2 percent of applicants receive an initial allowance, while 19.4 percent decide not to appeal an initial denial. The percentages of claimants that end up in the final allowance and final denial categories are 24.9 percent and 9.5 percent, respectively. For comparison, the bottom panel of the table displays equivalent quantities

Table 2.
Number and percent of sample observations, by adjudicative disability category, 1997–2004

Count and proportion	Initial		Final		Total
	Allowances	Denials not appealed	Allowances	Denials	
			10 percent random sample		
Number	213,851	89,796	115,112	43,819	462,578
Percent	46.23	19.41	24.88	9.47	99.99
			100 percent data file		
Number	2,319,171	1,053,375	1,265,000	513,805	5,151,351
Percent	45.02	20.45	24.56	9.97	100.00

SOURCE: Author's calculations based on a 10 percent sample of the DRF and Table 1.

NOTE: Values may not sum to 100 because of rounding.

corresponding to the full data set. The outcome proportions in the 10 percent random sample suggest an adequate approximation to the population of DI claimants over the 8-year period.[8]

Summary statistics of the explanatory variables used in my modeling effort appear in Table 3. Age at filing is the only continuous predictor. As illustrated in a later section of this article, the age profiles associated with the disability outcomes are highly nonlinear. In the models, I include both age and its square as a means to capture the nonlinearity. The mean age of all filers in the sample is about 50, but on average, claimants receiving an initial allowance tend to be 2 years older, while those in the final denials category have a mean age of less than 47. All else equal, it is expected that an increase in age would be positively associated with the likelihood of an initial allowance.

The models include binary indicators for sex (1 if male), for reapplication (1 if the claimant has applied to the DI program before), and for having zero earnings in the year before application (1 for zero earnings). Males comprise 52 percent of all filers in the sample, but make up 56 percent of claimants receiving an initial allowance. All else equal, it is expected that males would have a higher probability of an initial allowance. The two remaining indicator variables (reapplicants and claimants with zero earnings in the year before filing) are included because of their potential to serve as proxies for marginally qualified applicants, however imperfectly.

Following the DRF documentation, I use a 10-year window to classify an individual as having previously applied. That is, a new claimant is a person who is actually a first-time applicant or whose previous DI application dates back at least 10 years. About 17 percent of filers in the sample are reapplicants, compared with only 12 percent of those receiving an initial allowance. Notice how outcomes in the final adjudicative stages tend to have a higher share of claimants with a prior application history. Thus, it is expected that new applicants would have a higher likelihood of an initial allowance. Finally, the focus turns to a claimant's lack of earnings in the year before filing to identify those with the highest immediate financial incentive to apply. Throughout this study, such applicants are referred to as unemployed (Table 3). About 19.5 percent of claimants in the sample had zero earnings in the year before applying, compared with 24 percent and 28 percent of those in the initial denials not appealed and final denials categories, respectively. All else equal, it is anticipated that applicants with nonzero earnings in the year before filing would have a higher probability of an initial allowance.

The last explanatory variable used here is a derived field in the DRF, representing a discrete earnings index. The earnings index is constructed using the Department of Labor's official minimum wage and Social Security's Office of the Chief Actuary's national income averages. An applicant's individual earnings are compared with the minimum wage and the national income average in order to assign a numerical value (from 1–5) that indicates whether the claimant's earnings are below or above the national average. Among allowed claims, the index encompasses the 2nd through 6th years of earnings prior to

Table 3.
Summary statistics of explanatory variables (in percent)

Variable	Initial: Allowances	Initial: Denials not appealed	Final: Allowances	Final: Denials	Total (variable category)
Male	56.29	49.21	49.20	48.15	52.38
Reapplicant	11.82	18.62	22.09	23.41	16.79
Unemployed	15.13	24.21	20.72	27.63	19.47
Earnings					
Marginal	22.47	36.91	23.63	37.45	26.98
Low	25.50	29.46	29.96	28.88	27.70
Average	26.34	19.63	25.34	19.59	24.15
High	18.44	10.77	15.88	10.86	15.60
Very high	7.25	3.23	5.19	3.22	5.58
Age					
Mean	52.15	47.39	49.58	46.76	50.08
Standard deviation	10.10	10.80	8.52	9.31	10.03

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

the established date of disability onset. Among the denied claims, the earnings index comprises the 2nd through 6th years of earnings before the filing date. The rationale in choosing this time frame is based on a desire to avoid potential bias that is due to a sharp decline in earnings in the most recent years because of the gradual onset of disability. The earnings index categories are as follows:

1. Marginal earnings.
2. Low earnings—mean earnings exceed marginal earnings, up to 75 percent of the national average.
3. Average earnings—mean earnings fall between 75 percent and 125 percent of the national average.
4. High earnings—mean earnings fall between 125 percent and 200 percent of the national average.
5. Very high earnings—mean earnings above 200 percent of the national average.

While zero earnings in the year before filing (defined here as unemployed) reflects a claimant's immediate incentive to apply, the earnings index encompasses the future earnings potential that the applicant must renounce in order to receive DI benefits. Roughly, 27 percent of filers have marginal earnings, which tend to distribute more heavily among the denial categories (36.9 percent of initial denials not appealed and 37.5 percent of final denials). That trend reverses for average, high, and very high earners. For instance, 15.6 percent of claimants in the sample are high earners. However, among applicants receiving an initial or a final allowance, their shares are 18.4 percent and 15.9 percent, respectively. Meanwhile, the proportion of high-income filers in each of the initial denials not appealed and final denials categories is less than 11 percent. All else equal, it is anticipated that higher earnings would be associated with a higher probability of an initial allowance.

The Models

The Bayesian approach to inference embodies the idea of learning from experience, through which new evidence is integrated with existing knowledge. Given observed data, a researcher (classical or Bayesian) makes probabilistic assumptions about how that data were generated (the data distribution or data model). The model contains a number of unknown parameters and the goal is typically to reach statistical conclusions about their values. Bayesian statisticians include a second element to the model (the prior distribution), which reflects prior uncertainty about the parameter values. Those two elements are combined through a mechanism known as Bayes's theorem to derive the so-called posterior distribution. The *posterior probability distribution* results from conditioning on the observed sample and reflects how the information in the data modifies prior knowledge. Once available, it can be used to report point estimates of the parameters, construct credible intervals and regions of the parameter space associated with some posterior probability, and estimate the posterior predictive density associated with future observations.

The *prior probability distribution* (often called the prior) provides a formal mechanism to explicitly incorporate available nonsample information. The prior might be specified to accommodate the empirical evidence of previous studies or for purely economic or statistical theory considerations. It may also aim at simply reflecting the views of the researcher. These are examples of informative priors. On the other hand, diffuse or noninformative priors aim at representing a lack of prior knowledge, by minimizing the influence of the prior on the resulting posterior distribution. At any rate, when a large sample of observations is involved, the data density usually dominates the prior, so that the choice of prior is inconsequential in terms of the derived posterior inference.[9]

The Bayesian models estimated in this analysis closely follow the description and algorithmic implementation in Rossi, Allenby and McCulloch (2005). I estimate separate hierarchical multinomial logit models that cluster the claimant-level data into states and into diagnoses. Appendix Tables A-1 and A-2 present sample counts by disability outcome for the 181 primary impairments and 50 states, respectively. Following Congdon (2005), a hierarchical multinomial logit model is often defined by the nature of the individual-level explanatory variables entertained. In this application, all of the available predictors are invariant with respect to the adjudicative disability outcome. As a result, the specification becomes a pure multinomial logit model with category-specific parameters. The parameters for a baseline outcome are typically set to zero to avoid model indeterminacy. In all cases, final denials represent the baseline. Thus, for a particular cluster (a specific state or diagnosis) and a particular outcome (an initial allowance, an initial denial not appealed, or a final allowance), there is a distinct set of parameters associated with the following explanatory variables:

- An intercept.
- A binary variable taking the value of 1 if the individual has applied to the DI program before.

- A binary variable taking the value of 1 if male.
- A discrete earnings index taking values of 1 through 5.
- A binary variable taking the value of 1 if the individual had zero earnings in the year before applying.
- The applicant's age at filing.
- The square of the applicant's age at filing.

One way to think of a hierarchical model is as a compromise between two extreme solutions. On the one hand, I could disregard the state of origin and the primary diagnosis codes and estimate a multinomial logit model that pools all the claimants together. For comparison, estimates from such a model are provided. Alternatively, I could estimate a separate model for every state and every impairment. That approach would be problematic for those groups with few observations, which is the case for many of the individual impairments. Instead, the hierarchical version of the model can be seen as a set of multinomial logit processes that are linked together through a common distributional assumption. That is, the individual parameters are assumed to derive from a multivariate normal distribution (often referred to as the heterogeneity distribution), with unknown mean and covariance matrix. Estimates of the covariance matrix can be used to decompose outcome variation into its within-group and between-group components (see for instance, Raudenbush and Bryk (2002)). Moreover, unlike the nonhierarchical version of the multinomial logit model, my approach can accommodate the possibility of correlation between the groups, although not within the groups. Finally, one virtue of hierarchical models lies in their ability to diminish the influence of outlying observations. That property (often referred to as shrinkage) is desirable in circumstances where many of the clusters contain few observations. The result is usually more reasonable parameter estimates that are not skewed by the scarcity of data or the influence of outliers in specific groups.

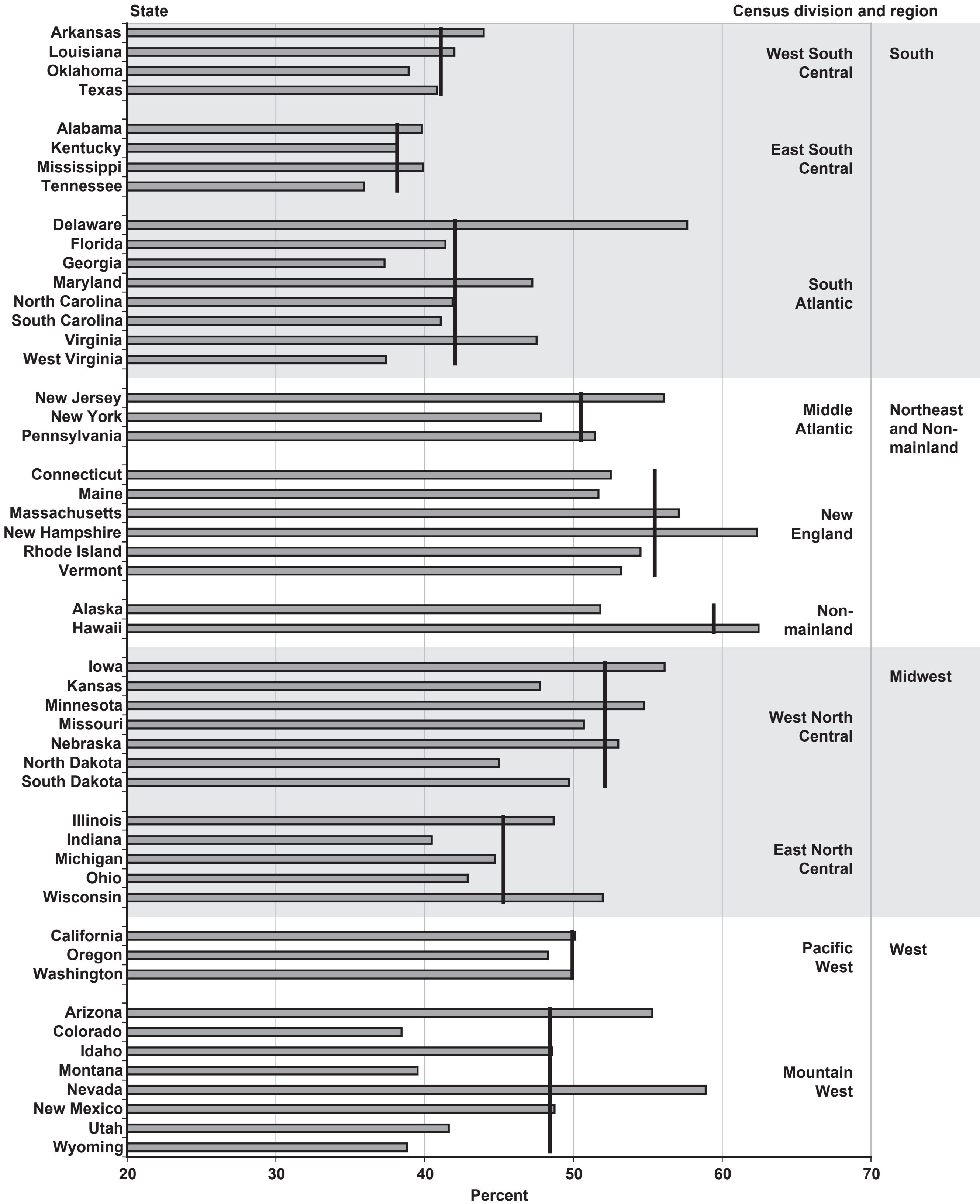
Once posterior estimates of the parameters are available, the models can be used to generate probability predictions.[10] Given specific values of the explanatory variables, three separate equations generate linear predictions for an initial allowance, an initial denial not appealed, and a final allowance (by default, the linear prediction for a final denial takes a 0 value). These linear predictions can be transformed into probabilities using standard formulae associated with the logit model. It is important to keep in mind the distinction between a linear prediction and a probability. For a given outcome (say an initial allowance), the linear predictions allow comparison of how all the clusters (the states or diagnoses) rank within that outcome. On the other hand, the probability that the i-th applicant in the j-th group falls into say the initial allowance category is computed using the linear predictions for all four adjudicative disability outcomes combined. Thus, within a given cluster, the estimated probabilities of an initial allowance, an initial denial not appealed, a final allowance, and a final denial add to 100 percent, as they track the observed proportions in the data sample.

State Variation

The disability outcomes in the sample for all 50 states are listed in Appendix Table A-2. In terms of sample size, California contributes 10.1 percent of total applicants, followed by New York, Florida, and Texas. These four states combined account for more than a quarter of all claimants. At the other end of the spectrum, Alaska comprises a mere 0.12 percent of the total observations (552), followed by Wyoming, North Dakota, and South Dakota. The graphs in Chart 1 display initial allowance rates by state, grouped according to the Census Bureau regions and divisions. The black vertical lines denote the overall initial allowance rate for a particular division, with the horizontal bars corresponding to each individual state. For geographical reasons, I place Alaska and Hawaii in the Nonmainland category, although technically, those two states are counted as part of the Pacific-West division.

In terms of initial allowance rates, the four states with the lowest values are southern states: Tennessee (35.9 percent), Georgia (37.3 percent), West Virginia (37.4 percent), and Kentucky (38.1 percent). On the other hand, Hawaii leads with the highest initial allowance rate at 62.5 percent, followed by New Hampshire (62.3 percent), Nevada (58.9 percent), and Delaware (57.7 percent). Thus, the range of state variation in initial allowances (the difference between Hawaii with the highest initial allowance rate and Tennessee with the lowest rate) is roughly 25 percentage points. Chart 1 does not appear to reveal any clear-cut geographical patterns other than perhaps the contrast between the South and New England. Specifically, the three divisions with the lowest initial allowance rates are the southern ones (West South Central, East South Central, and South Atlantic). Clearly, Delaware and to a lesser extent Maryland and Virginia appear to be outliers in the South Atlantic division and more at

Chart 1.
Percentage of initial allowances, by state and Census division and region



SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

NOTE: The black vertical lines indicate the percentage for each Census division.

home in the Middle Atlantic division. Overall, however, it is fair to say that southern states tend to have low initial allowance rates. New England, on the other hand, is the Census division with the highest allowance rate.

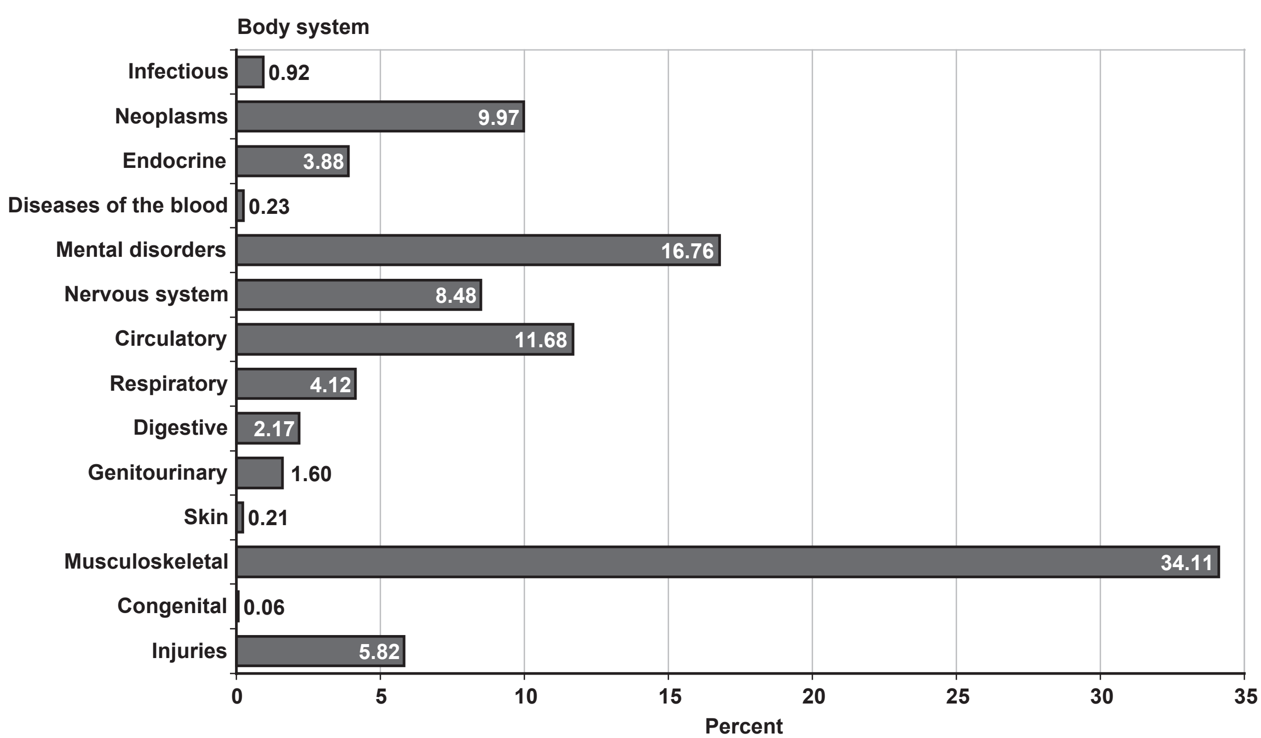
Diagnosis Variation

SSA maintains a classification of impairments that identify the medical conditions on which disability-related claims are based. Since 1985, the coding of primary and secondary diagnoses has approximately followed the *International Classification of Diseases: 9th Revision* (ICD-9) taxonomy. Appendix Table A-1 summarizes the disability outcomes for 181 medical impairments, which are grouped into 14 body systems.[11] Notice that I employ the body system for descriptive purposes only, as a means of grouping individual diagnoses. To this end, each impairment is uniquely matched to a single body group, following the description in the *SSA Program Data User's Manual* (Panis and others 2000).

The primary diagnosis field in the data is generally based on the latest Form SSA-831 at the DDS level, but will be assigned based on an alternative source if that field is incomplete. There is evidence that on appeal, some claimants will be evaluated on the basis of a different primary diagnosis. That may occur for a number of reasons. Typically an adjudicator designates the primary impairment at the time of the decision, based on the medical evidence. However, many disability claims allege multiple impairments. Moreover, impairments may worsen and new diagnoses develop over time. As a result, additional medical evidence introduced on appeal can lead an adjudicator to change the primary impairment. Unfortunately, the DRF does not identify changes in the primary diagnosis throughout the adjudicative process. Such events are not accommodated in this analysis. An audit report from Social Security's Office of the Inspector General (SSA 2010) found that a switch in the primary diagnosis was common for three of the four impairments most likely to be denied at the initial level and allowed at the hearing level in the 2004–2006 period. These three impairments (diabetes mellitus; osteoarthrosis and allied disorders; and muscle, ligament, and fascia disorders) are prone to worsen over time and affect other body systems.[12]

Chart 2 displays the percentage of claimants in each body system for the entire sample. Musculoskeletal impairments account for 34 percent of the diagnoses,

Chart 2.
Percentage of claimants, by body system



SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

followed by mental disorders with 17 percent. Those two body systems combined make up slightly over half of all observations. Circulatory diseases and neoplasms represent 12 percent and 10 percent of all outcomes, respectively. The nervous system and sense organs category comprises 8 percent of the impairments, while injuries make up 6 percent. Both the respiratory and the endocrine, nutritional, and metabolic body systems account for about 4 percent of claimants each. Likewise, each of the digestive and genitourinary body systems represents 2 percent of all diagnoses. Infectious and parasitic diseases contribute almost 1 percent of the observations. Finally, the remaining body groups (congenital anomalies and both diseases of the skin and subcutaneous tissue and blood and blood forming organs) represent well below 1 percent of cases combined.

A cursory look at Appendix Table A-1 reveals that one or a few primary diagnoses codes may sometimes account for the bulk of diagnoses within a body system. The tabulation below highlights selected cases. For example, disorders of the back and osteoarthrosis represent 56 percent and 21 percent of all musculoskeletal impairments, respectively, while affective and mood disorders make up more than half of the mental diagnoses. Diabetes and obesity respectively contribute 63 percent and 31 percent of claimants to the endocrine, nutritional, and metabolic body system. Four types of cancers (lung, breast, colon, and genital organs) comprise over 50 percent of the neoplasms.[13] Similarly, symptomatic HIV infections are more than half of all infectious and parasitic disorders. Chronic liver disease and cirrhosis accounts for 56 percent of digestive impairments, while about 67 percent of respiratory ailments involve chronic pulmonary insufficiency. Finally, 85 percent of the genitourinary impairments are chronic renal failure, which explains the high initial allowance rate of this body system.

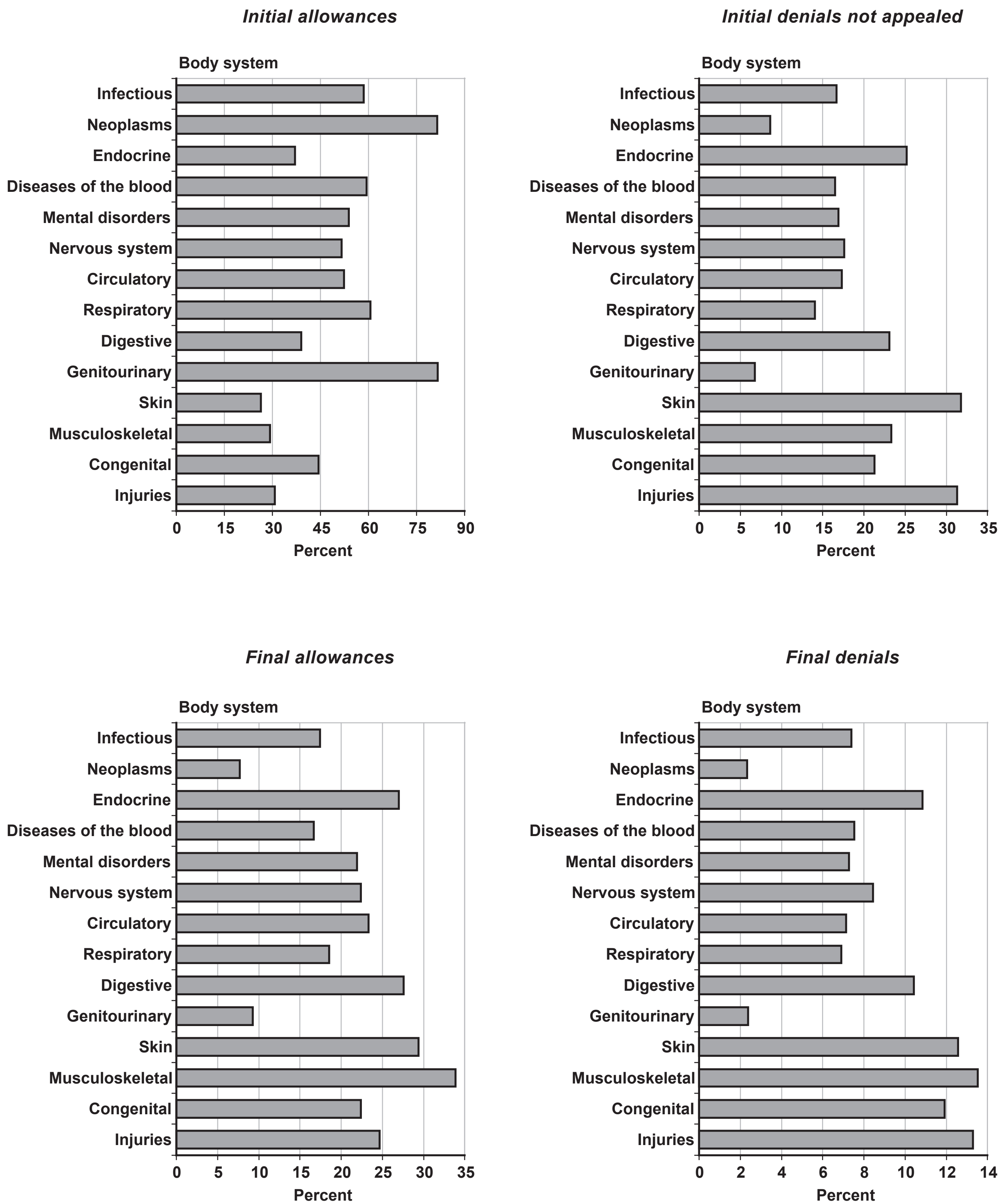
Impairment	Percent
Musculoskeletal	
Disorders of the back—discogenic and degenerative	55.7
Osteoarthrosis and allied disorders	20.8
Mental	
Affective/mood disorders	55.7
Neoplastic	
Malignant cancers of the—	
Trachea, bronchus, or lung	19.0
Breast	15.5
Colon, rectum, or anus	10.0
Genital organs	9.2
Respiratory	
Chronic pulmonary insufficiency	66.7
Endocrine, nutritional, and metabolic	
Diabetes	62.6
Obesity and other hyperalimentation disorders	30.7
Digestive	
Chronic liver disease and cirrhosis	55.7
Genitourinary	
Chronic renal failure	84.9
Infectious and parasitic	
Symptomatic HIV infections	52.8

There is huge variation in disability outcomes by primary diagnosis. Chart 3 illustrates the proportion of decisions that correspond to each body system. The overall proportion of initial allowances in the sample is 46.2 percent (Table 2). However, over 80 percent of genitourinary and neoplastic impairments receive an initial allowance, while the share drops to 26.3 percent for skin disorders and to about 30 percent for musculoskeletal diagnoses. Thus, the range of variation in initial allowances among the body systems is roughly 55 percentage points. In general, the genitourinary and neoplastic body systems have the highest initial rates of allowance, exceeding any other group by at least 20 percentage points. As a result, those two groups also have the lowest proportions of initial denials not appealed, final allowances, and final denials. Applicants with injuries and skin impairments appear most likely not to appeal an initial denial, with about 31 percent of the outcomes. Musculoskeletal diagnoses have the highest proportion of final allowances, with about 34 percent of the outcomes, followed by skin disorders. In addition to injuries, however, musculoskeletal and skin impairments also exhibit the highest rates of final denials.

Mortality Variation

One source of concern regarding the categorization of outcomes in this analysis is a potential biasing effect that is due to death. Specifically, claimants with an initial denial could die before having a chance to appeal. Our DRF sample identifies an applicant's date of death over the 11-year period from 1997 through 2007. It is of course impossible to determine from the data which deaths occurred as a direct result of the underlying disability impairment. Nevertheless, this information is used to compute raw death rates (adjusted neither by age or sex) over the period in question. For the different body systems, Table 4 shows the proportion of applicants in every adjudicative outcome that passed away. About 17 percent of all claimants died during this period. However, while 28.4 percent of

Chart 3.
Percentage of adjudicative disability categories, by body system



SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

the applicants in the initial allowance category died, only about 7 percent of claimants who did not appeal an initial denial did not survive to 2007. Among those, two-thirds passed away at least 3 years after their application. Consequently, the potential fraction of applicants who died before having the chance to appeal would be too marginal to affect this analysis in any material way.

Deaths occurred more frequently among the most medically serious diagnoses. In terms of all outcomes, the body system with the lowest rate of mortality during the 11-year period is musculoskeletal, which is followed by injuries, mental disorders, and skin impairments. The diagnostic groups with the highest proportion of deceased claimants are neoplasms, followed by genitourinary impairments, diseases of the blood and blood forming organs, respiratory diagnoses, and digestive disorders. Given the DI program's goal to serve claimants in greater need more expeditiously, it is reassuring to see that the proportion of deceased claimants in every single body system is highest among those initially allowed and second highest for filers in the final allowance category.

It is also worth recalling that disability in the DI program is defined on the basis of long-term inability to work. As a result, death proportions and initial allowance rates are not expected to always go hand in hand. For instance, 82 percent of claimants with a neoplasm disorder who receive an initial allowance die within the 11-year period under consideration. For corresponding applicants with a genitourinary disorder (85 percent of whom have a diagnosis of chronic renal failure), mortality is lower (39 percent). Nevertheless, both body systems have similar initial allowance rates of roughly 81 percent. Standard treatments for those two impairments (such as chemotherapy and dialysis) likely pose equally severe barriers to work, even if one kind of diagnosis is much more deadly in the short run.

Age Variation

Another relevant factor of variation in disability adjudicative outcomes is age. Three important characteristics are identified in the data:

1. The proportion of outcomes by single year of age is both highly nonlinear and pretty regular from one year to the next.
2. There are distinct patterns at ages 50 and 55, which represent threshold points in the vocational grid.
3. There is an age-62 effect that results from an influx of early retirement applicants. As pointed out by Leonesio, Vaughan, and Wixon (2003), it is a common procedure at SSA field offices to compare the potential benefits to which an applicant is entitled under more than one program. What this means in practice is that early retirees with health problems often apply concurrently for retirement and disability benefits.

Table 4.
Percentage of applicant deaths, by adjudicative disability category and body system, 1997–2007

Body system	Initial		Final		Total (claimant deaths in the period)
	Allowances	Denials not appealed	Allowances	Denials	
All	28.37	6.94	8.75	5.62	17.17
Infectious	30.55	10.06	14.88	7.03	22.66
Neoplasms	82.27	21.88	38.78	16.54	72.21
Endocrine	23.84	11.82	14.54	10.48	16.86
Diseases of the blood	43.50	9.83	18.86	8.86	31.24
Mental disorders	8.49	5.41	6.74	5.25	7.35
Nervous system	16.03	5.40	7.59	5.58	11.39
Circulatory	25.30	12.49	14.31	10.05	19.44
Respiratory	37.83	11.42	15.08	9.29	27.95
Digestive	47.50	11.67	18.36	10.77	27.37
Genitourinary	39.20	9.80	20.26	10.80	34.79
Skin	15.69	5.19	8.42	4.10	8.76
Musculoskeletal	7.72	4.01	4.97	3.68	5.37
Congenital	18.70	8.47	9.68	3.03	12.64
Injuries	12.33	4.39	5.61	3.91	7.07

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

Chart 4 displays the number of claimants for each adjudicative disability outcome by single year of age (18–65). Because the focus here is on workers covered by the DI program, the total number of applicants at the youngest ages represents a tiny fraction of the sample (239 claimants at age 18 out of more than 462,000 observations). At ages 30–49, the rate at which applicants join the initial allowance category is fairly constant, but increases sharply by age 50 (top graph on the left). There are also noticeable spikes at ages 55 and 62, the latter representing a peak with over 14,000 observations. On the other hand, the number of claimants initially denied who decide not to appeal rises at a fairly constant rate up until about age 42, but levels off subsequently. The most remarkable feature in the top right graph of Chart 4 is the huge spike at age 62. The number of applicants at age 62 in this category totals more than twice that of filers at ages 61 or 63. This suggests that a substantial portion of concurrent early retirement and DI applicants receive an initial denial and decide against filing an appeal. The graph on final allowances (bottom left) shows visible spikes at ages 50 and 55, while final denials experience a jump at age 62 (bottom right).

The proportion of outcomes (rates) by single year of age is shown in Chart 5. The thin discontinued lines in the chart denote the age profiles for each individual year from 1997 through 2004, while the continuous thick line corresponds to the full 8 years of data combined. The proportion of initial allowances by age displays a distinct convex "u-shape," while initial denials not appealed, final allowances, and final denials roughly follow a concave profile in the form of an "inverted-u." These patterns exhibit a great deal of regularity from one year to the next.

For the youngest claimants, the initial allowance rate is very high, ranging from 60 to 70 percent at ages 18–23 (top graph on the left). Then, the rate declines rapidly, reaching 34 percent by age 30, where it remains stable in the low-to-mid 30 percent range until age 49. The subsequent increase resembles a piece-wise linear function with discontinuities at ages 50 and 55 and a dip at the early retirement age. The rate of initial denials not appealed (top graph on the right) rises from about 20 percent at age 20 to its peak of 35.5 percent by age 27. It steadily declines from this point forward, reaching its lowest value of 11 percent at age 59. As retirement nears, the rate increases again, with the early retiree effect inducing a sizeable jump at age 62. The final allowance rate (bottom graph on the left) rises steadily to its peak of 34 percent at age 50, declining rapidly afterwards. Finally, the rate of final denials (bottom graph on the right) hovers below 15 percent at ages 32–48, declining to about 5 percent by age 55.

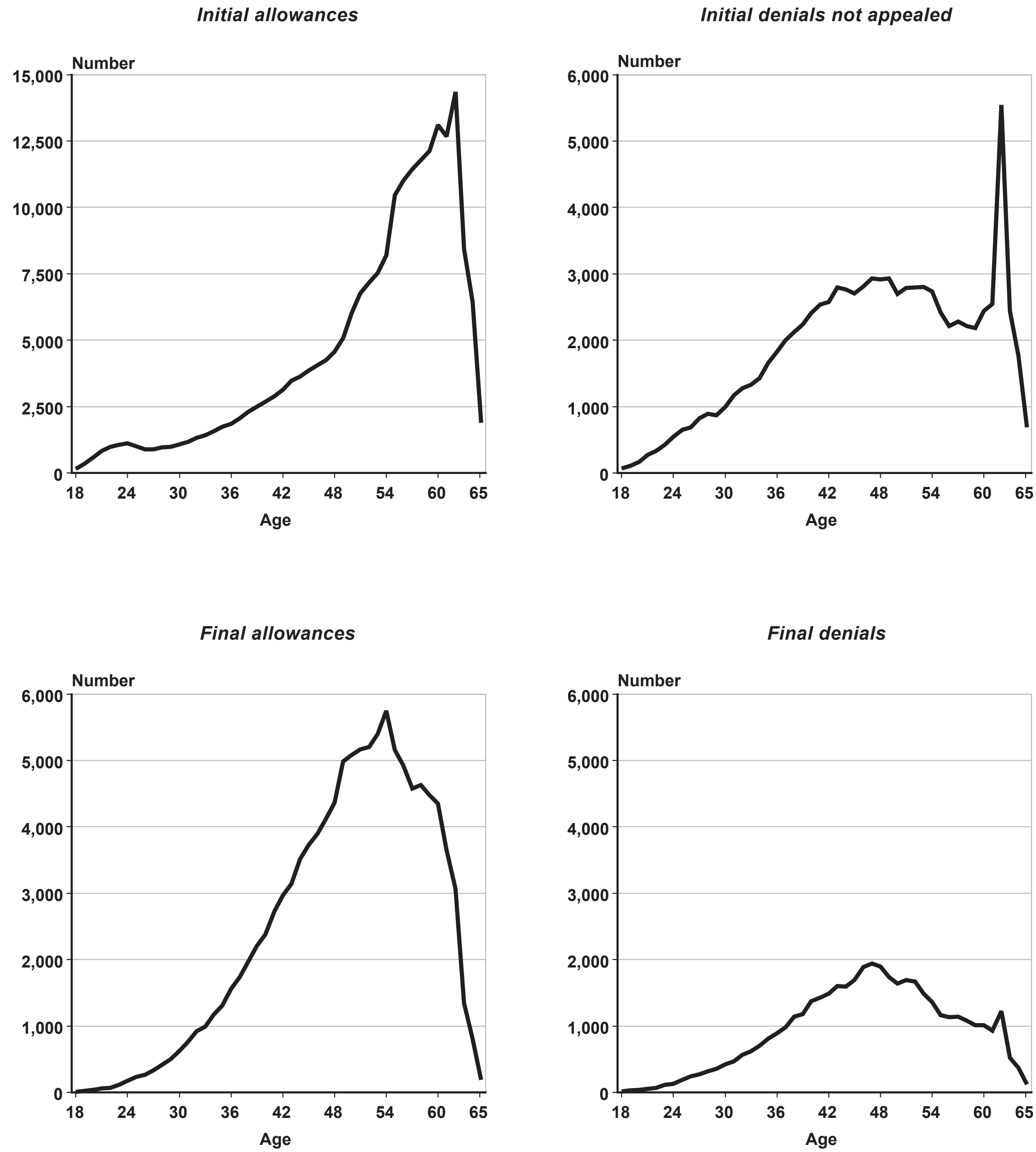
One interesting aspect of the age profiles is their nonlinearity. Specifically, the convex shape in the proportion of initial allowances might appear at odds with the notion that age is a reasonable proxy for health. Beyond some threshold age range, it is reasonable to expect the initial allowance rate to rise. After all, the increasing prevalence of serious age-related disabilities and less stringent vocational standards of the program are bound to push allowance rates upward. But what explains the high initial allowance rates for claimants at a very young age? One plausible answer is that the high allowance rates are driven by the impairment severity of a tiny number of applicants from an otherwise very healthy pool of workers. In addition, the contributory requirements of the DI program could be creating a bottleneck effect, with young disabled workers waiting to reach insured status. A look at the diagnostic makeup of claimants by age reveals some insights.

Chart 6 displays the distribution of claimants for the most common body systems by single year of age. About 60 percent of the small fraction of applicants aged 18–23 receive a mental diagnosis. Because mental impairments tend to have a very early onset, they indeed dominate the composition of claimants until about age 30. From age 31 forward, musculoskeletal impairments become the most common diagnosis. On the other hand, the share of mental impairments declines steadily with age. By ages 55 and 57, circulatory disorders and neoplasms surpass mental impairments to respectively become the second and third leading groups of diagnoses.

Inferential Results

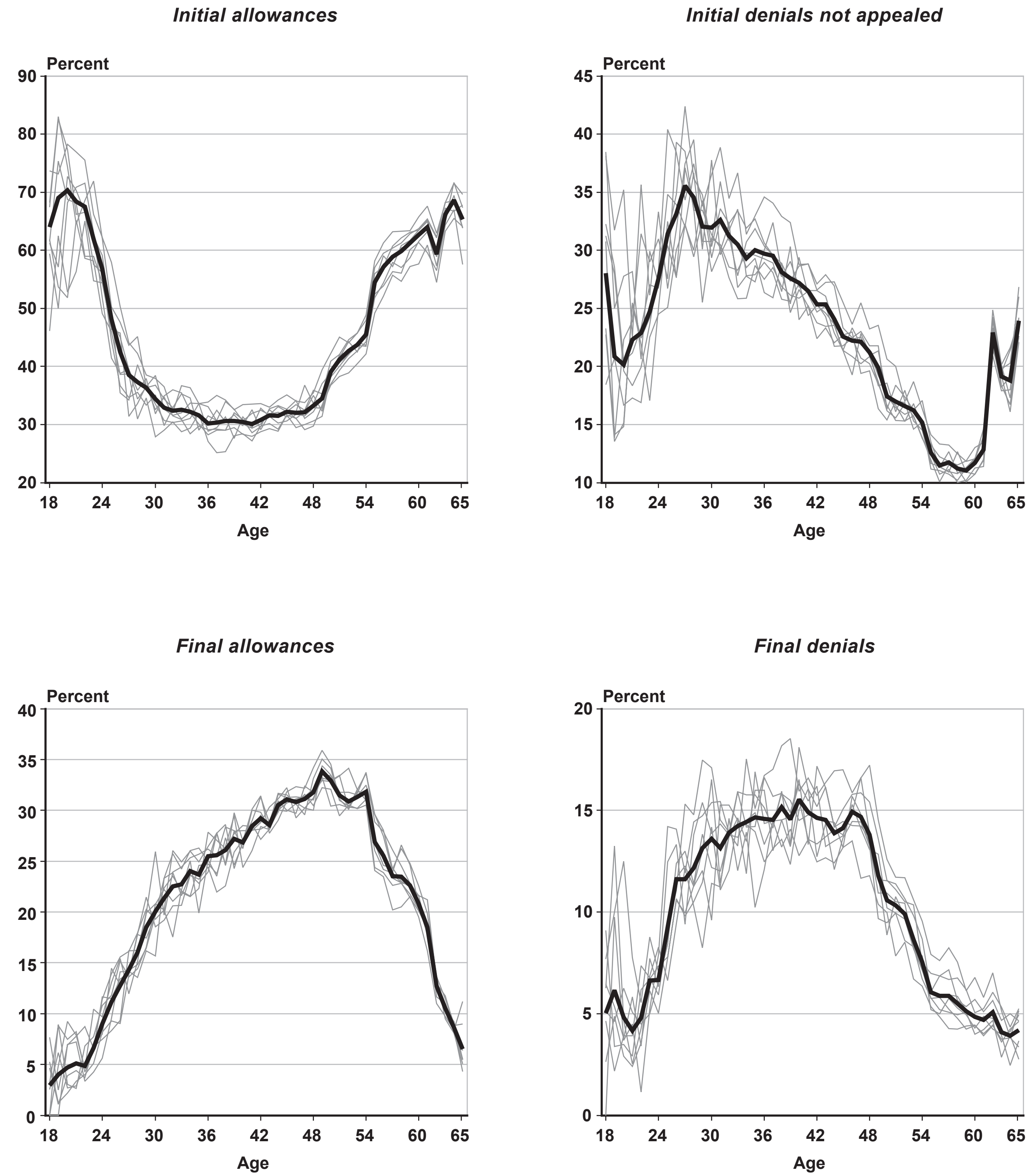
For each hierarchical structure (claimants nested by state or diagnosis), two model specifications are contemplated. Each model is estimated initially with no explanatory variables other than intercepts. The intercepts-only specification is useful to apportion unconditional data variance between hierarchical levels. It also provides a benchmark lower bound to goodness-of-fit criteria, which can be used for comparison purposes. The second specification entertains the previously described individual-level predictors. In addition, estimates are provided for a pooled or nonhierarchical model that does not entertain any grouping of the data.

Chart 4.
Number of claimants, by adjudicative disability category and single year of age



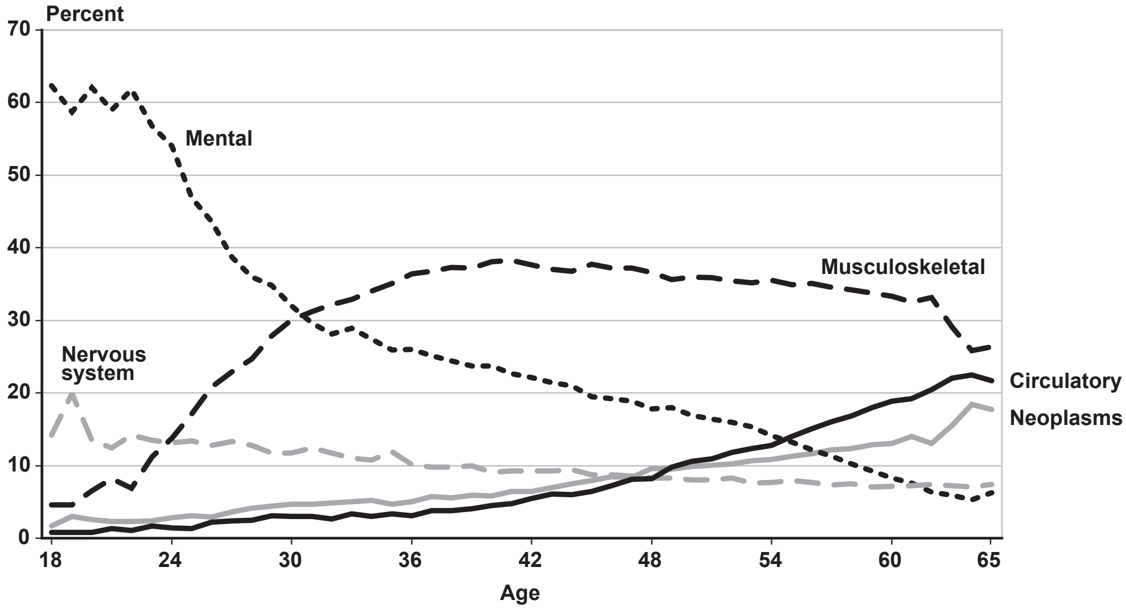
SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

Chart 5.
Percentage of adjudicative disability categories, by single year of age



SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

Chart 6.
Percentage of claimants, by selected body systems and single year of age



SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

Next, I consider two different metrics for goodness-of-fit assessment. One measure that is particularly convenient in the context of Bayesian hierarchical models is the deviance information criterion (DIC), proposed by Spiegelhalter and others (2002). The DIC can be seen as the Bayesian analogous to the classical Akaike information criterion. It incorporates cross-validation and penalizes excess complexity. When comparing multiple specifications, the smaller the DIC value, the better the model's fit. DIC estimates are presented in the following tabulation. Additionally, I compute the percentage of observations correctly predicted by each model, shown in Table 5. In this case, an observed outcome is treated as a correct prediction if its estimated posterior mean probability is higher than the mean classification probabilities of the three other remaining outcomes.

Model specification	DIC estimate
Intercepts only	
Pooled	1,151,155.30
State	1,140,108.40
Diagnosis	1,038,875.60
Individual-level inputs	
Pooled	1,093,989.10
State	1,080,995.40
Diagnosis	980,212.70

Both measures of model fit provide a consistent picture. First, for a given set of variables, there is an unequivocal advantage in grouping claimants by state rather than pooling them together and in grouping them by impairment rather than clustering them by state. Consider for instance the top entry in Table 5, which corresponds to the intercepts-only pooled multinomial logit specification. As there are no explanatory variables, the estimated probability of any observation within a category is simply the sample proportion. All claimants are predicted to receive an initial allowance because this is the outcome that occurs most often. As a result, all of the initial allowances, but none of the other outcomes, are correctly categorized. This provides a lower predictive bound of 46.23 percent of the decisions correctly classified.

One way to think of a model with only intercepts is as a naive classification rule. In a hierarchical context, all individual outcomes within say a state or a diagnosis are predicted to be equal to the disability category with the highest sample proportion for that state or diagnosis. In grouping claimants by state, the intercepts-only model variant achieves some very modest gains relative to the pooled specification (46.26 percent). On the other hand, prediction improves more significantly if claimants are clustered by diagnosis (51.45 percent). When claimant-level explanatory

Table 5.
Percentage of observations correctly predicted, by model and adjudicative disability category

Model	Initial		Final		Total (correctly categorized)
	Allowances	Denials not appealed	Allowances	Denials	
			Intercepts only		
Pooled	100.00	0	0	0	46.23
State	97.18	0	5.38	0	46.26
Diagnosis	83.68	18.06	37.18	0	51.45
			Individual-level inputs		
Pooled	90.80	6.35	17.07	0	47.46
State	85.89	9.43	27.95	0.03	48.50
Diagnosis	87.84	24.80	39.50	0.20	55.27

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

variables are accommodated, the hierarchical diagnosis model can accurately classify 55.27 percent of the observations. The DIC estimates result in a similar ranking of the models.

A second conclusion can be drawn from Table 5. Notice how the diagnosis model with only intercepts correctly predicts a larger share of observations (51.45 percent) than the state model with claimant-level explanatory variables (48.50 percent). The same conclusion is reached when comparing the DIC estimates in the tabulation on the previous page (1,038,875 versus 1,080,995). This suggests that the primary diagnosis codes carry greater predictive ability than all other explanatory variables that are entertained combined. To put it differently, grouping a sample of claimants by diagnosis alone (the naive classification rule implied by an intercept-only model) will predict the adjudicative disability decision outcomes more accurately than knowing everything else, including age, sex, state of origin, application history, earnings history, and employment status in the year before filing. This finding is hardly unexpected, considering the role medical impairments play in the disability determination process. However, the result suggests that the full range of primary diagnosis codes (which are often overlooked for the purpose of research) is a crucial piece of information among the limited set of useful variables typically available from administrative data extracts.

Average Effects

The top portion of Table 6 presents posterior means and standard deviations of the regression coefficients in the pooled multinomial logit model.[14] The bottom part of the table displays estimates corresponding to the so-called average effects of the hierarchical diagnosis model. These parameters represent the mean of the distribution of the diagnosis-specific coefficients (that is, the estimated means of the multivariate normal heterogeneity distribution). For both models (pooled and hierarchical), the estimates tend to have similar signs and magnitudes, although as expected, the standard deviations are much higher in the hierarchical version of the process.

Given a particular observation and model, three equations yield continuous linear predictions of an initial allowance, an initial denial not appealed, and a final allowance. Those linear predictions are defined in reference to the benchmark category of final denials, which has a zero linear prediction by design. All else equal and relative to an initial denial, the sign of the estimated coefficients implies the following effects at the claimant level:

- The linear prediction of an initial allowance: (1) increases for males and higher earners, and (2) decreases for unemployed applicants and claimants who have applied before.
- The linear prediction of an initial denial not appealed: (1) increases for males; and (2) decreases for higher earners, unemployed applicants, and claimants who have applied before.
- The linear prediction of a final allowance: (1) increases for higher earners and claimants who have applied before, and (2) decreases for males and unemployed applicants.

At the individual level, the estimated effects for the explanatory variables match my a priori expectations. The results also appear consistent with research

Table 6.
Posterior parameter means and standard deviations, by adjudicative disability category

Variable	Initial allowances: Mean	Initial allowances: Standard deviation	Initial denials not appealed: Mean	Initial denials not appealed: Standard deviation	Final allowances: Mean	Final allowances: Standard deviation
			Pooled multinomial logit			
Intercept	1.252581	0.007014	0.568303	0.007492	1.060417	0.007195
Reapplicant	-0.427993	0.013526	-0.207347	0.014696	0.143670	0.013884
Male	0.121928	0.011101	0.026245	0.012352	-0.086577	0.011496
Earnings	0.236173	0.005119	-0.013485	0.005773	0.215522	0.005406
Unemployed	-0.614464	0.013845	-0.159045	0.013977	-0.292570	0.013669
Age	0.085465	0.000775	0.031423	0.000837	0.031753	0.000817
Age^2	0.004253	0.000051	0.002366	0.000055	-0.000036	0.000058
			Hierarchical diagnosis multinomial logit (average effects)			
Intercept	1.682253	0.131225	0.698521	0.046508	1.213439	0.057412
Reapplicant	-0.363732	0.060715	-0.198089	0.061695	0.202028	0.061481
Male	0.200885	0.054438	0.107004	0.052963	-0.069526	0.054504
Earnings	0.242488	0.040234	-0.029794	0.039516	0.220765	0.041297
Unemployed	-0.655286	0.064591	-0.213411	0.062811	-0.332018	0.064286
Age	0.081470	0.030319	0.027600	0.030594	0.016404	0.029204
Age^2	0.011028	0.027359	0.001629	0.027164	0.002343	0.027872

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

by Rupp (2012), who also used claimant-level data. Specifically, Rupp's "fixed-effects" binary logit model for initial determinations yielded qualitatively similar conclusions about the impact of sex and unemployment on the initial allowance rate. Of course, there are substantial differences in the two modeling approaches. Rupp (2012) used the time-varying state unemployment rates, while I do not control for year-effects and instead define unemployment at the individual level (as having zero earnings in the year prior to application). All else equal, the higher the earnings category, the higher the opportunity cost of filing for DI benefits, which may explain the positive association I find between earnings and the predictions of both an initial and a final allowance. Meanwhile, a history of previous applications shows a negative impact on the likelihood of an initial allowance, but a positive impact on the likelihood of a final allowance. In addition, I find that reapplicants are more likely to appeal an initial denial.

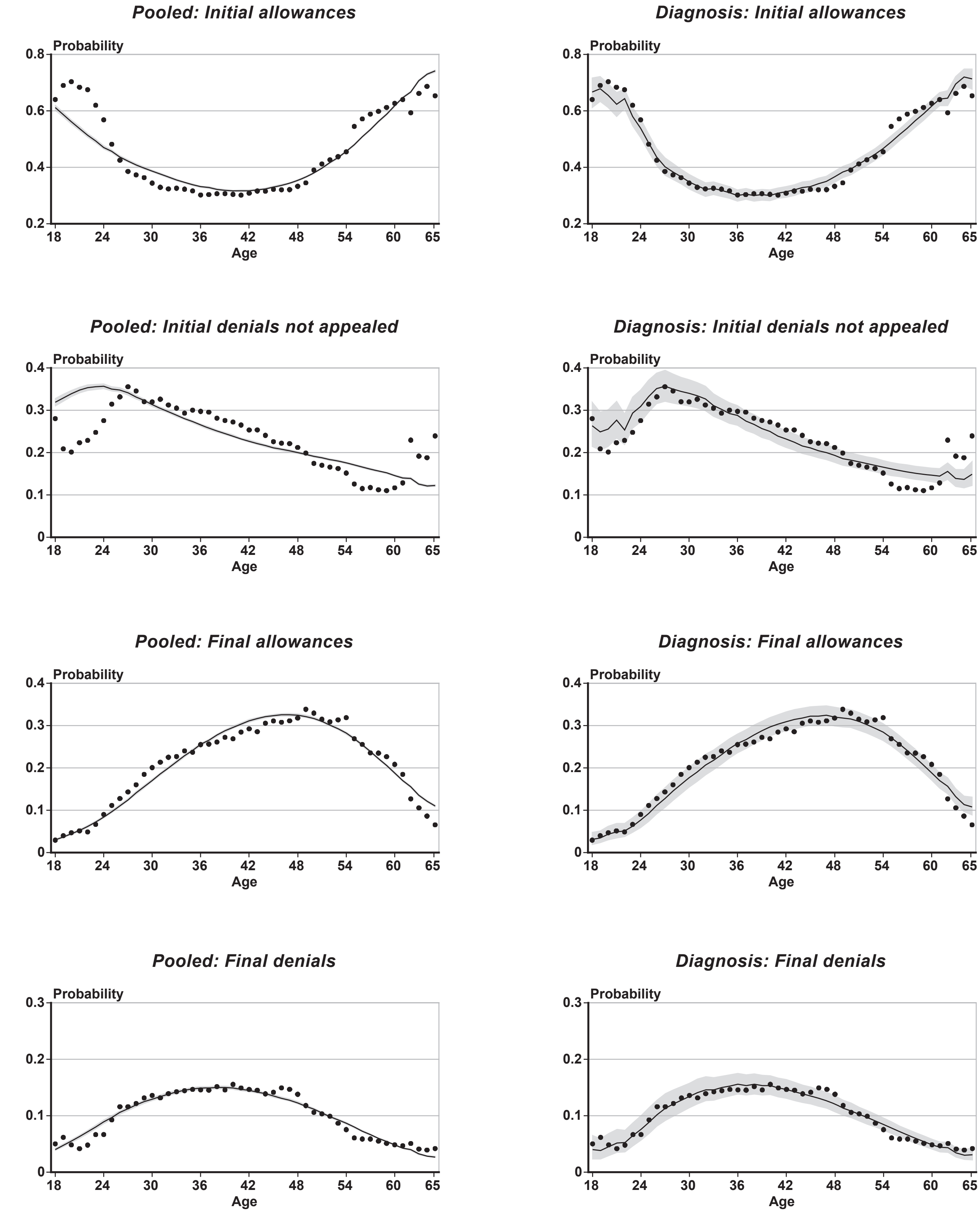
The interpretation of the parameters associated with age is less tractable because of the fact that those parameters represent the coefficients of a quadratic polynomial. Aggregate point and interval probability predictions for each outcome by single year of age are presented in Chart 7. Those predictions are obtained by averaging over the estimated probabilities of all the claimants in the sample who are the same age. The shaded areas in the graphs represent 90 percent posterior credible intervals (in other words, intervals containing 90 percent posterior probability). The thin dark lines along the intervals correspond to the posterior mean of each prediction. In addition, the solid dots show the actual proportions observed in the sample.

In general, it appears that the square term for age does a reasonably good job at capturing the nonlinear shape of the age profiles. The left and right columns of graphs in Chart 7 correspond to the pooled and hierarchical diagnosis models, respectively. The interval estimates for the pooled specification seem inadequately narrow, seriously underrepresenting uncertainty, as they miss most of the actual proportions. The point and interval predictions for the hierarchical diagnosis process clearly provide an improvement in fit. This is particularly evident in both the greater width of the intervals and at the youngest ages, where the shape of the age profiles is defined by relatively small numbers of claimants with a predominance of mental impairments.

Variance Decomposition

One issue of particular interest in this analysis is variance decomposition; that is, the portion of total variation in outcomes that the models attribute to

Chart 7.
Aggregate point and interval probability predictions for each adjudicative disability category, by single year of age



SOURCE: Author's calculations based on a 10 percent random sample of the DRF and model estimates.

the groups rather than the claimants. The top panel of Table 7 presents posterior means and standard deviations of between-group variances for the specifications with intercepts only. Consider for instance the first entry in the table, which corresponds to an initial allowance in the state hierarchical specification. The model has 50 intercept parameters per equation, each representing a state's mean linear prediction of an initial allowance. The posterior mean of the variance among those predictions is 0.22. Likewise, the between-state variance estimate for the linear prediction of an initial denial not appealed is 0.16.

In a similar fashion, the middle panel of Table 7 shows between-group variances corresponding to the models with claimant-level explanatory variables. Now the intercepts represent mean linear predictions of the outcomes when the explanatory variables take their average values in the sample.[15] Thus, the adjusted mean linear prediction of an initial allowance has a between-state variance of 0.59. Likewise, the variance of the mean-adjusted predictions for an initial denial not appealed between the states is 0.51.

One pattern emerges from the estimates in Table 7. For a given specification, the between-state variances corresponding to the prediction of all three outcomes are small and close in magnitude to one another. On the other hand, things are quite different when claimants are grouped by their impairments. In particular, variation in the prediction of an initial allowance between the diagnoses is very large (2.6 for the model with only intercepts and 2.8 for the variant with individual explanatory variables). Those magnitudes dwarf the variances associated with the other adjudicative categories (initial denials not appealed and final allowances). The implication is one of considerable heterogeneity in the prediction of an initial allowance among the impairments. This is of course consistent with the description of the data, where some primary diagnosis codes have initial allowance rates of over 95 percent, while others are close to zero.

In hierarchical models, total data variance is the sum of the within-group and the between-group variances. A useful statistic of variance decomposition is the intraclass correlation coefficient (ICC), which measures the proportion of variance in the outcomes between the groups. A value close to zero indicates a good deal of homogeneity between the clusters, so that most of the data variance can be attributed to individual-level variation within the groups. Conversely, an ICC close to 100 percent suggests a high degree of between-group heterogeneity, which implicitly favors a hierarchical modeling structure.

The bottom panel of Table 7 displays estimated ICC values.[16] On average, only about 6.2 percent of total variance in initial allowances can be attributed to differences between the states. Most of the observed heterogeneity in initial allowances (over 90 percent) seems to be due to disparities among claimants within the states. The decomposition suggests that applicants within any given state can be very heterogeneous in

Table 7.
Posterior estimates of group-level variances and ICCs, by adjudicative disability category

Disability outcome	State: Mean	State: Standard deviation	Diagnosis: Mean	Diagnosis: Standard deviation
		Between-group variances: Intercepts only		
Initial allowances	0.219	0.044	2.587	0.286
Initial denials not appealed	0.160	0.032	0.122	0.016
Final allowances	0.180	0.036	0.269	0.035
		Between-group variances: Individual-level inputs		
Initial allowances	0.594	0.120	2.824	0.338
Initial denials not appealed	0.514	0.104	0.274	0.034
Final allowances	0.543	0.108	0.409	0.052
		ICCs (percent)		
Initial allowances	6.22	1.17	43.89	2.70
Initial denials not appealed	4.65	0.90	3.59	0.46
Final allowances	5.17	0.98	7.56	0.91

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

NOTE: ICC = intraclass correlation coefficient.

their disability characteristics. In fact, once claimants are grouped by primary diagnosis, a large portion of variation previously attributed to the individuals can now be explained by the differences between the impairments. About 44 percent of total variation in initial allowances is attributed to the different diagnosis groups. These results do not extend to the other outcomes (initial denials not appealed and final allowances), where group-level heterogeneity does not exceed 10 percent of total variance.

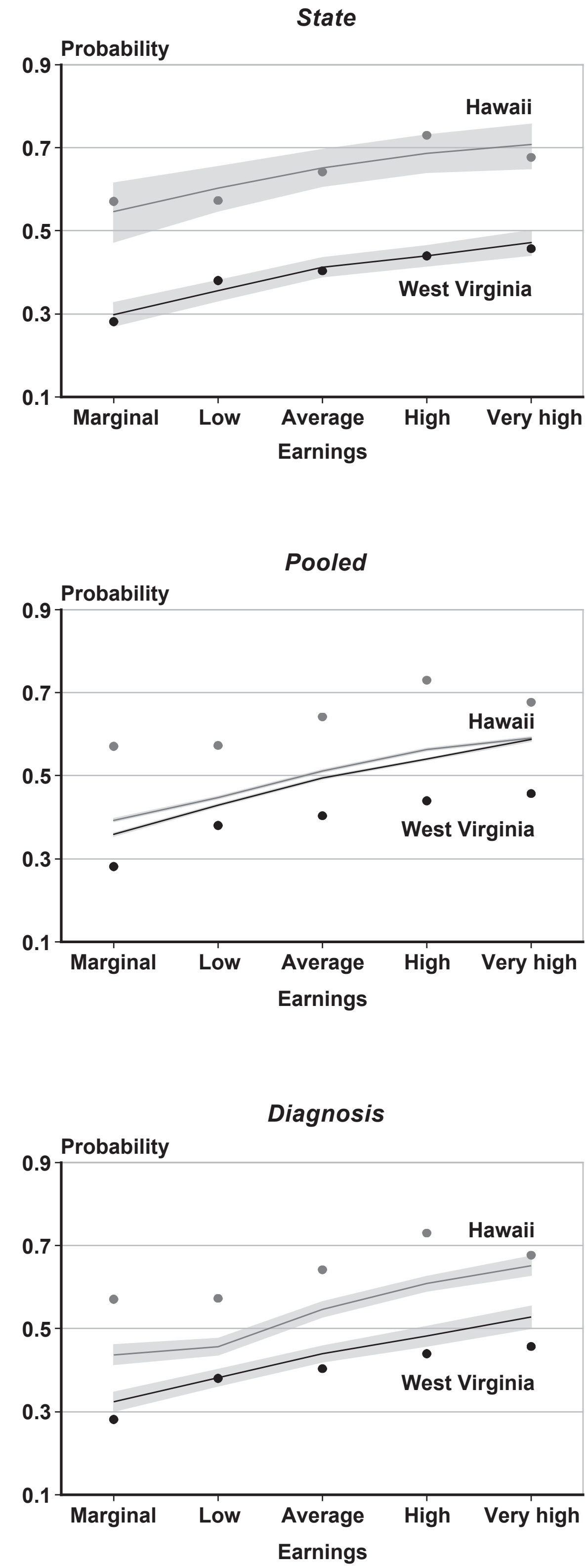
One of the implications of the ICC estimates is that the primary diagnoses can account for a great deal of the observed variation in initial allowances among claimants. To the extent that it is possible, parallels are drawn between the findings in this article and those in Rupp (2012). Fixed-effects models are not designed to apportion variance into between-group and within-group sources. Rupp (2012, Table 9) looked at the decomposition of overall variation in initial allowance rates across states by three sources. For adult DI-only claimants, the state fixed-effects accounted for 52 percent of the variation, while the year fixed-effects and the demographic and diagnostic characteristics of claimants contributed 14 percent and 10 percent of variation, respectively. The large size of the state fixed-effects in Rupp's article suggested that long-term unique differences among the states were substantial. That might seem at odds with this article's finding of small between-state, but large within-state variation in the outcomes. Notice, however, that the hierarchical state model here tracks with a great deal of accuracy the four adjudicative outcomes for each one of the states. This is by design because the model accommodates state-specific parameters. In other words, the hierarchical state model does a much better job at predicting the observed allowance and denial rates by state than does the hierarchical diagnosis model. Nevertheless, as the DIC tabulation and Table 5 confirm, the hierarchical diagnosis model unquestionably fits the overall data much better. First, it yields a significantly smaller DIC estimate. Second, for all claimants, it correctly predicts a higher share of each of the four adjudicative outcomes than does the state model.

The results in Rupp (2012) hinted at the diagnostic mix playing a role (although a small one), in explaining state heterogeneity in initial allowance rates.[17] The findings here (values not shown) are consistent with that view, in that the diagnostic mix is not a major factor at accurately predicting initial allowance rates in most states, except in some cases, despite the fact that state variation in the composition of impairments is substantial in the sample under study. For instance, the proportion of musculoskeletal diagnoses ranges from 27 percent in Hawaii to 42.9 percent in Montana. Mental disorders comprise 26.9 percent of the diagnoses in New Hampshire, but only 12.1 percent of those in Arkansas. Neoplasms vary from 13.6 percent in Iowa to 6.3 percent in West Virginia. Mississippi has the highest composition of circulatory diagnoses at 15.8 percent, while Idaho has the lowest at 7.1 percent. For the nervous system and sense organs group, Colorado has a proportion of diagnoses (12.3 percent) that is three times the size of that corresponding to Vermont. Injuries also vary from 2.5 percent in South Dakota to 10.6 percent in West Virginia. Coe and others (2011) cited substantial variation in age-adjusted mortality rates by state and even greater variation in self-reported disability.

In the context of my modeling effort, one way to further illustrate state heterogeneity in disability outcomes is through a specific example. Chart 8 provides a comparison between the states of Hawaii and West Virginia. The graphs display point and interval probability predictions (90 percent posterior probability) of an initial allowance as a function of earnings for both states. Hawaii exhibits the highest initial allowance rate in the sample at 62.5 percent. In addition, it also happens to have the lowest proportion of musculoskeletal impairments of any state. By contrast, West Virginia has the third-lowest initial allowance rate (37.4 percent) and incidentally, the lowest proportion of neoplasms and the highest share of injuries among the 50 states.

The top graph in Chart 8 corresponds to the hierarchical state model, which by design, accurately reproduces the observed state proportions. Notice that Hawaii has a smaller number of observations than West Virginia (Appendix Table A-2), resulting in state-specific parameter estimates with greater variance (and as a result, a wider probability interval). The middle graph in Chart 8 presents the predictions associated with the pooled model. In this case, there is a wide gap between observed and predicted outcomes. Over all claimants, Hawaii and West Virginia differ in their proportion of initial allowances by about 25 percentage points (see Chart 1). Instead, the pooled model predicts a mean gap of about 3 percentage points, despite the fact that the predictions take into account the different mix of characteristics between the applicant populations in the two states (age, sex, employment status, application history, and earnings history).

Chart 8.
Aggregate point and interval probability predictions for an initial allowance, by earnings: Hawaii compared with West Virginia



SOURCE: Author's calculations based on a 10 percent random sample of the DRF and model estimates.

The graph at the bottom of Chart 8 shows the probability predictions resulting from the hierarchical diagnosis model. This specification incorporates the same individual-level predictors as the pooled multinomial logit model. The only difference, of course, is that claimants are grouped according to their impairments. Relative to the observed proportions, the diagnosis model slightly overpredicts the probabilities corresponding to West Virginia, but significantly underpredicts the probabilities associated with Hawaii. On average, the predicted gap in the probability of an initial allowance between the two states is 11 percentage points. In other words, discrepancies in claimant-level characteristics (differences in the impairment mix specifically) seem to account for a little less than half of the observed difference in the initial allowance rate between these two states. This result, however, does not generalize to comparisons among other states.

Correlation Across Outcomes

Table 8 presents posterior estimates of the correlation between the disability adjudicative outcomes. The top panel of the table corresponds to the intercepts-only specification, while the bottom panel comprises the estimates for the models with claimant-level predictors. For example, the mean correlation between the average linear predictions of an initial allowance and an initial denial not appealed among the 50 states is 0.25. Likewise, the mean correlation between those two outcomes among the 181 primary diagnosis codes is 0.31. When the individual explanatory variables are included in the models, the corresponding correlation for the adjusted linear prediction of an initial allowance and an initial denial not appealed is 0.1 among the states and 0.13 among the impairments.

A look at Table 8 reveals that after controlling for individual-level predictors, the correlations in the state hierarchical models are small in magnitude and statistically insignificant. However, when claimants are grouped by diagnosis, there is very high statistically significant positive correlation between the linear predictions of an initial and a final allowance. For instance, with only intercepts, the posterior mean correlation among the impairments is 0.74. After controlling for claimant-level explanatory variables, a mean estimate of 0.56 is obtained. To the best of my knowledge, the finding of high significant positive correlation when impairments are used as a criterion for grouping claimants has never been reported in the literature. The finding is important for several reasons. First, it indicates that the zero correlation property implicit in a pure multinomial logit model (the so-called independence from irrelevant alternatives property) is an unrealistic restriction. More generally, any effort to model the adjudicative process using the impairments should accommodate this pattern in the data.

My classification of claimants roughly corresponds to a two-stage adjudication (decisions at the DDS level versus decisions made mostly at the hearing level or above). In this context, the estimation results suggest a substantial degree of dependence between the two adjudicative outcomes. Across the impairments, the high positive correlation between the predictions of an initial and a final allowance is important for a second reason. Normatively speaking, the more disabling a diagnosis, the greater the linear predictions of both an initial and a final allowance should be, relative to less disabling impairments. In this very narrow sense, the correlation result here appears to suggest a degree of consistency within the adjudicative process.

Consider the top graph on the left in Chart 9, which plots posterior means of the intercepts for the 181 primary diagnosis codes corresponding to the model with claimant-level predictors. Those coefficients represent adjusted mean linear predictions of an initial denial not appealed and a final allowance. There is no apparent relationship between the two outcomes, as a statistically insignificant mean correlation estimate of 0.12 bears out in Table 8. Transforming the linear predictions into actual probabilities results in the top graph on the right. Unlike the linear predictions, the probabilities show an upward trend. Impairments that have a higher classification probability of an initial denial not appealed also tend to have a higher probability of a final allowance.

The bottom-left graph in Chart 9 plots the relationship between the linear predictions of an initial and a final allowance for each of the impairments. In this case, the mean correlation is 0.56 (shown in Table 8). However, the corresponding probabilities in the bottom-right graph indicate the opposite effect (negative correlation). In other words, diagnoses that have a higher classification probability of an initial allowance tend to have a lower classification probability of a final allowance. The reason for the correlation inversion has to do with the fact that the probability of an outcome is a nonlinear function of the linear prediction of all the possible outcomes. As the linear prediction of an initial allowance dominates the magnitude of the other predictions, the classification probabilities of an initial denial not appealed, a final allowance, and a final denial can only decline.

The implications of high positive correlation between the linear predictions of an initial and a final allowance (bottom-left graph in Chart 9) can be further clarified with a somewhat extreme example involving the two impairments that are presented in Chart 10. The most common diagnosis in the musculoskeletal body system is a disorder of the back (discogenic and degenerative). The proportions in the entire sample of initial and final allowances for that impairment are about 23 percent and 38 percent, respectively. On the other hand, based on its effect on mortality alone, a highly disabling diagnosis is lung cancer (malignant neoplasm of the trachea, bronchus, or lung). In this case, 94 percent of the decisions result in an initial allowance, while only 3 percent of the outcomes represent a final allowance.

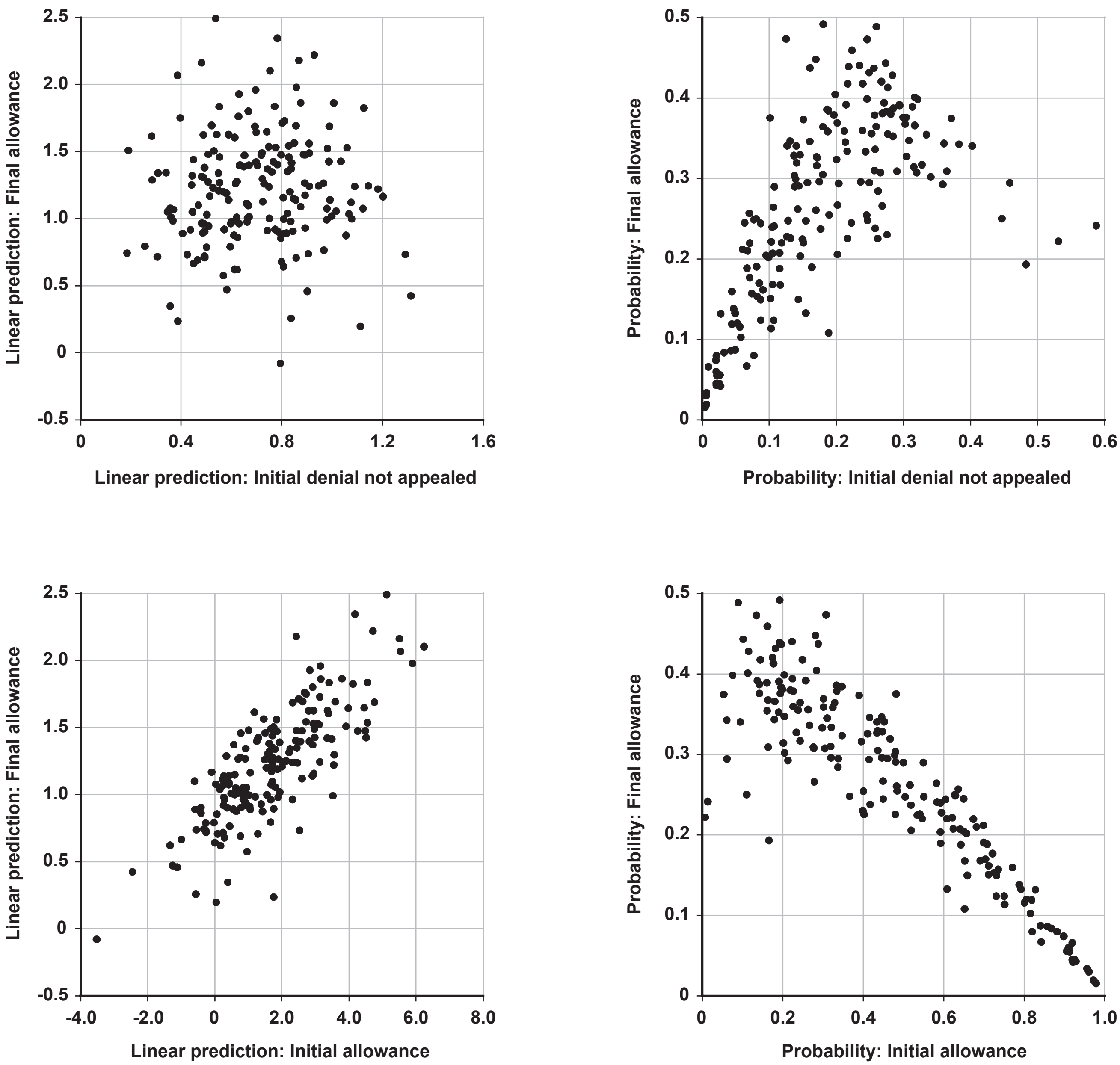
Table 8.
Posterior correlations, by model specification

Correlation sequence of disability outcome	State: Mean	State: Standard deviation	Diagnosis: Mean	Diagnosis: Standard deviation
		Intercepts only		
Initial allowance—initial denial not appealed	0.249	0.129	0.307	0.087
Initial allowance—final allowance	0.063	0.136	0.737	0.041
Initial denial not appealed—final allowance	0.015	0.135	0.177	0.092
		Individual-level inputs		
Initial allowance—initial denial not appealed	0.100	0.133	0.125	0.096
Initial allowance—final allowance	0.048	0.135	0.561	0.064
Initial denial not appealed—final allowance	0.006	0.134	0.119	0.087

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

Suppose two claimants were identical in all measured characteristics (having the sample mean features), except one was diagnosed with lung cancer and the other had a back disorder. Linear predictions for those two claimants as a function of earnings appear on the left (top and bottom) graphs of Chart 10. Notice in particular how the predictions of an initial and a final allowance for the claimant with lung cancer exceed the predictions corresponding to the applicant with a back disorder. By contrast, the two graphs on the right side of the chart display point and interval probability predictions (90 percent posterior probability), which closely follow the observed sample proportions. For any outcome different from an initial allowance, the classification probabilities of lung cancer lie well below the probabilities of a disorder of the back. This, of course, is due to the extremely high probability of an initial allowance

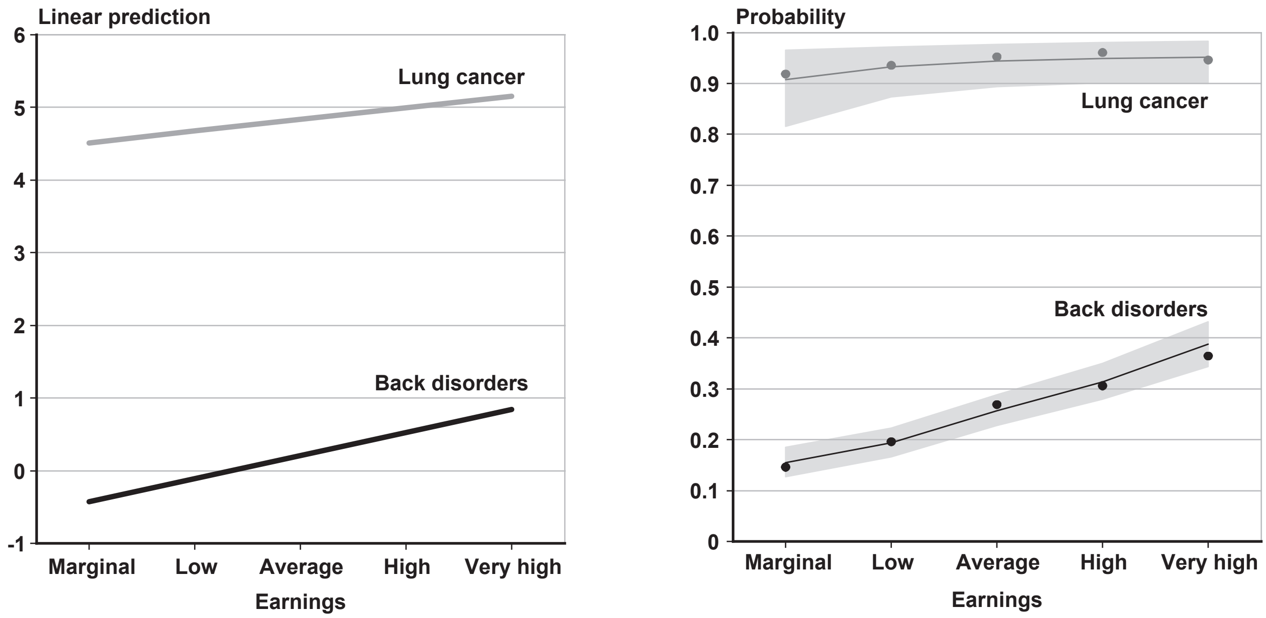
Chart 9.
Linear predictions compared with probabilities in the diagnosis model



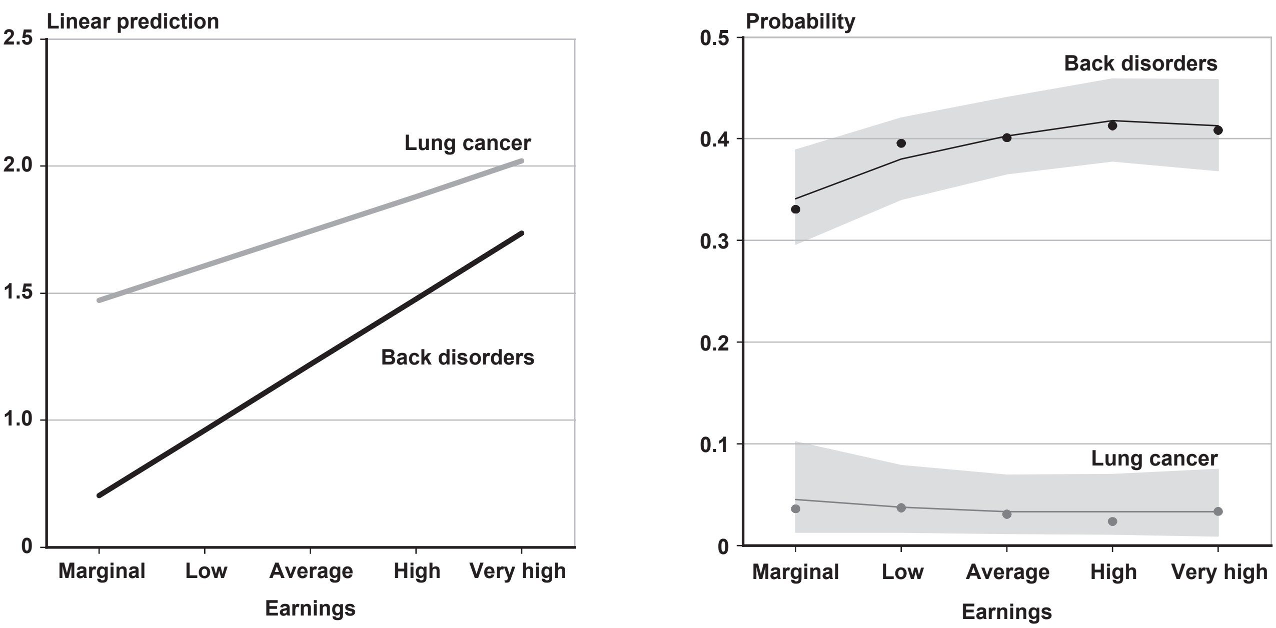
SOURCE: Author's calculations based on a 10 percent random sample of the DRF and model estimates.

Chart 10.
Lung cancer versus disorders of the back, by earnings: Linear predictions compared with probabilities

Initial allowances



Final allowances



SOURCE: Author's calculations based on a 10 percent random sample of the DRF and model estimates.

associated with a diagnosis of lung cancer in the first place.

In the two-impairment (lung cancer/back disorder) example, a significant fraction of claimants with back disorders are initially denied, but eventually allowed. Yet, claimants with lung cancer have a higher prediction of both an initial and a final allowance. Put differently, it is simply not the case that ALJs are favoring applicants with back disorders over those with lung cancer. Whether it is at the DDS or at the hearing level or above, lung cancer is determined to be a more disabling diagnosis than a back disorder. In general, the high positive correlation implies that in going from an initial to a final allowance, decision makers are largely preserving the ordinal ranking of impairments (a finding that is only evident when looking at the linear predictions and not the probabilities).

One might be tempted to conclude that this correlation finding provides evidence that decision makers are uniformly adhering to SSA's disability guidelines at the various adjudicative levels. However, other possible explanations cannot be ruled out. For example, Keiser (2010) hinted at evidence of a policy feedback mechanism, where knowledge of ALJ reversal rates affected the self-reported initial allowance rate of DDS examiners.[18] If there was a feedback effect, it could also flow in either direction (from the DDS to the Office of Disability Adjudication and Review (ODAR) and vice versa), or from both directions simultaneously. The bottom line is that it is important not to overreach when it comes to interpreting my results. The positive correlation between the predictions of an initial and a final allowance could be potentially explained by a feedback effect, where decision makers at the two stages are influenced by each other's ranking of impairments. Nevertheless, whether a feedback mechanism or adherence to the guidelines explains the positive correlation, the result implies some degree of consistency.

Conclusion

This article explores the roles that primary diagnoses and state of origin play in explaining observed heterogeneity in disability outcomes by adjudicative stage. Disability determinations are separated into four mutually exclusive categories: (1) initial allowances, (2) initial denials not appealed, (3) final allowances, and (4) final denials. The main findings are as follows:

- The primary diagnosis codes carry greater predictive ability for placing claimants into adjudicative categories than all other explanatory variables that are entertained combined. Knowing the impairments of a sample of applicants yields more accurate classification probabilities than knowing their age, sex, state of origin, earnings, employment status in the year before filing, and application history combined.
- The prediction of an initial allowance (1) increases for males and higher earners, and (2) decreases for unemployed applicants and claimants who have applied before.
- The prediction of an initial denial not appealed (1) increases for males; and (2) decreases for higher earners, unemployed applicants, and claimants who have applied before.
- The prediction of a final allowance (1) increases for higher earners and claimants who have applied before, and (2) decreases for males and unemployed applicants.
- As a function of single year of age, the initial allowance rate has a *u-shape* defined at very young ages by small numbers of claimants with a predominance of mental impairments. A quadratic polynomial seems to reproduce the age profiles accurately.
- When claimants are grouped by state, variance decomposition suggests that most of the variation in outcomes is driven by individual-level heterogeneity within the states. On the other hand, almost half of the variation in initial allowances can be attributed to the various primary diagnoses. In some cases, the different mix of impairments in the population of claimants may explain a significant portion of the difference in initial allowances between two states. Still, a great deal of variation in outcomes remains unaccounted for by the models, particularly when it comes to identifying final denials.
- When applicants are grouped by diagnosis, there is high positive correlation between the predictions of an initial and a final allowance. To the best of my knowledge, that finding has never been documented in the literature. Impairments that are considered to be more disabling at the DDS level tend to also be considered more disabling at the hearing level or above. In other words, when moving from an initial to a final allowance, the severity ranking of the diagnoses is preserved to a good extent.

Table A-1.
Sample distribution, by adjudicative disability category, body system, and primary diagnosis

Body system and primary diagnosis	Initial		Final		Total
	Allowances	Denials not appealed	Allowances	Denials	
Infectious/parasitic diseases	2,478	706	739	313	4,236
Pulmonary tuberculosis	(X)	13	(X)	(X)	27
Symptomatic HIV	1,559	298	283	98	2,238
Asymptomatic HIV	30	186	130	80	426
Neurosyphilis	19	(X)	(X)	(X)	36
Mycobacterial, other chronic infections	32	18	(X)	(X)	71
Other infectious and parasitic disorders	83	45	34	14	176
Late effects of acute poliomyelitis	568	50	115	31	764
Neoplasms	37,526	3,968	3,533	1,070	46,097
Malignant neoplasm of tongue	254	21	24	9	308
Malignant neoplasm of salivary glands	(X)	(X)	(X)	(X)	21
Malignant neoplasm of esophagus	1,123	(X)	36	(X)	1,179
Malignant neoplasm of stomach	641	(X)	24	(X)	687
Malignant neoplasm of small intestine	144	(X)	13	(X)	176
Malignant neoplasm of colon or rectum	3,528	514	435	126	4,603
Malignant neoplasm of liver	1,667	10	37	(X)	1,718
Malignant neoplasm of gallbladder	139	(X)	(X)	(X)	148
Malignant neoplasm of pancreas	1,357	(X)	24	(X)	1,394
Malignant neoplasm of digestive system	176	(X)	(X)	(X)	196
Malignant neoplasm of trachea or lung	8,249	161	281	50	8,741
Malignant neoplasm of pleura	332	(X)	(X)	(X)	347
Malignant neoplasm of heart	(X)	(X)	(X)	(X)	30
Malignant neoplasm of bone and cartilage	445	(X)	41	(X)	525
Malignant neoplasm of connective tissue	198	30	(X)	(X)	256
Malignant melanoma of skin	801	(X)	26	(X)	857
Other malignant neoplasm of skin	50	15	(X)	(X)	79
Malignant neoplasm of breast	4,717	1,370	731	345	7,163
Kaposi's sarcoma	(X)	(X)	(X)	(X)	(X)
Malignant neoplasm of bladder	451	65	57	12	585
Malignant neoplasm of kidney	977	64	63	23	1,127
Malignant neoplasm of eye	(X)	(X)	(X)	(X)	11
Malignant neoplasm of brain	2,507	55	111	21	2,694
Malignant neoplasm of nervous system	(X)	(X)	(X)	(X)	13
Malignant neoplasm of thyroid gland	87	38	33	13	171
Malignant neoplasm of endocrine glands	34	(X)	(X)	(X)	44
Malignant neoplasm of other sites (head, neck)	1,383	165	222	49	1,819
Secondary malignant neoplasms	232	(X)	(X)	(X)	244
Malignant neoplasm of unspecified site	47	(X)	(X)	(X)	58
Lymphoma	1,769	494	431	137	2,831
Multiple myeloma	900	45	123	12	1,080
Leukemias	1,626	89	128	25	1,868
Benign neoplasm of brain	430	159	208	73	870
Neoplasm of uncertain behavior	(X)	(X)	(X)	(X)	15
Neoplasm of unspecified/unknown nature	(X)	(X)	(X)	(X)	(X)
Malignant neoplasm of genital organs	3,191	520	395	123	4,229

Continued

Table A-1.
Sample distribution, by adjudicative disability category, body system, and primary diagnosis—*Continued*

Body system and primary diagnosis	Initial		Final		Total
	Allowances	Denials not appealed	Allowances	Denials	
Endocrine, nutritional, and metabolic	6,635	4,517	4,842	1,947	17,941
All disorders of thyroid	42	146	129	67	384
Diabetes mellitus	3,014	3,490	3,345	1,387	11,236
All disorders of parathyroid gland	(X)	(X)	(X)	(X)	(X)
All disorders of pituitary gland	(X)	(X)	15	(X)	28
All disorders of adrenal glands	(X)	(X)	(X)	(X)	22
Malnutrition (weight loss)	113	(X)	32	(X)	164
Disorders of plasma protein metabolism	(X)	(X)	(X)	(X)	(X)
Gout	65	75	75	37	252
Disorders of metabolism (cystic fibrosis)	85	(X)	11	(X)	112
Obesity and other hyperalimentation	3,229	725	1,140	416	5,510
Disorders of the immune mechanism	77	38	89	18	222
Diseases of the blood	623	173	175	79	1,050
Deficiency anemias	48	23	26	14	111
Hereditary hemolytic anemias	143	35	27	12	217
Aplastic anemia	152	(X)	21	(X)	184
Other anemias	148	53	39	15	255
Coagulation defects	(X)	(X)	(X)	(X)	28
Purpura and other hemorrhagic conditions	14	14	(X)	(X)	47
Other diseases of blood-forming organs	109	36	40	23	208
Mental disorders	41,770	13,117	17,007	5,641	77,535
Organic mental disorders	8,024	740	1,878	308	10,950
Schizophrenic, paranoid, psychotic disorders	3,963	650	665	186	5,464
Affective/mood disorders	19,678	8,466	11,290	3,768	43,202
Autistic disorders	75	(X)	(X)	(X)	89
Anxiety disorders	3,817	1,477	2,096	736	8,126
Personality disorders	457	280	177	119	1,033
Substance addiction (alcohol)	(X)	439	(X)	186	777
Substance addiction (drugs)	(X)	218	(X)	64	342
Somatoform disorders	216	61	162	32	471
Eating and tic disorders	(X)	(X)	(X)	(X)	(X)
Attention deficit disorder	44	32	11	19	106
Learning disorder	54	103	17	21	195
Mental retardation	5,079	335	311	81	5,806
Borderline intellectual functioning	354	310	184	119	967
Nervous system and sense organs	20,239	6,891	8,773	3,313	39,216
Cerebral degenerations	36	(X)	(X)	(X)	48
Brain atrophy	713	97	163	37	1,010
Parkinson’s disease	1,315	137	300	50	1,802
Anterior horn cell disease	690	(X)	(X)	(X)	740
Other diseases of spinal cord	763	43	110	18	934
Disorders of autonomous nervous system	155	67	107	33	362
Multiple sclerosis	3,543	588	1,612	311	6,054
Cerebral palsy	549	68	70	29	716
Epilepsy	727	1,290	901	592	3,510
Migraine	349	446	488	217	1,500
Other neurological conditions	1,365	888	1,135	485	3,873
Carpal tunnel syndrome	117	174	213	103	607
Diabetic and other peripheral neuropathy	2,478	554	1,186	292	4,510

Continued

Table A-1.
Sample distribution, by adjudicative disability category, body system, and primary diagnosis—*Continued*

Body system and primary diagnosis	Initial		Final		Total
	Allowances	Denials not appealed	Allowances	Denials	
Nervous system and sense organs (cont.)					
Myoneural disorders	430	260	376	159	1,225
Muscular dystrophies	532	65	162	44	803
Retinal detachments and defects	207	103	78	42	430
Other retina disorders	644	151	212	51	1,058
Glaucoma	200	126	94	65	485
Cataract	61	99	44	26	230
Visual disturbances	437	400	326	148	1,311
Blindness and low vision	2,838	554	552	226	4,170
Cardiac transplantation	62	(X)	(X)	(X)	75
Disorders of eye movements	(X)	(X)	(X)	(X)	13
Disorders of vestibular system	284	194	299	122	899
Other disorders of ear	38	127	58	54	277
Deafness	1,704	439	229	202	2,574
Circulatory	28,256	9,336	12,593	3,852	54,037
Rheumatic fever with heart involvement	(X)	(X)	(X)	(X)	(X)
Diseases of aortic valve	297	192	221	84	794
Other rheumatic heart disease	70	(X)	25	(X)	124
Essential hypertension	412	1,706	1,192	696	4,006
Hypertensive vascular disease	538	467	453	172	1,630
Hypertensive vascular and renal disease	14	(X)	(X)	(X)	30
Acute myocardial infarction	566	435	385	119	1,505
Angina without ischemic heart disease	126	106	113	37	382
Chronic ischemic heart disease	7,977	3,132	5,107	1,395	17,611
Chronic pulmonary heart disease	378	42	68	14	502
Valvular heart disease/other defects	229	169	217	81	696
Cardiomyopathy	2,514	554	1,016	264	4,348
Cardiac dysrhythmias	412	282	363	141	1,198
Heart failure	2,972	480	737	153	4,342
Late effects of cerebrovascular disease	7,786	984	1,478	371	10,619
Aortic aneurysm	201	72	100	26	399
Peripheral vascular (arterial) disease	2,373	254	550	96	3,273
Periarteritis nodosa, allied conditions	50	(X)	(X)	(X)	71
Disease of capillaries	(X)	(X)	(X)	(X)	(X)
Phlebitis and thrombophlebitis	106	72	79	38	295
Varicose veins of lower extremities	292	70	80	29	471
Other diseases of circulatory system	942	280	381	123	1,726
Respiratory	11,539	2,671	3,528	1,313	19,051
Chronic bronchitis	41	53	60	29	183
Emphysema	890	150	195	52	1,287
Asthma	786	1,148	996	564	3,494
Bronchiectasis	33	15	(X)	(X)	66
Chronic pulmonary insufficiency	9,271	1,014	1,898	509	12,692
Asbestosis	43	(X)	39	(X)	106
Pneumoconiosis	(X)	(X)	10	(X)	20
Other diseases of the respiratory system	471	270	318	144	1,203

Continued

Table A-1.
Sample distribution, by adjudicative disability category, body system, and primary diagnosis—*Continued*

Body system and primary diagnosis	Initial		Final		Total
	Allowances	Denials not appealed	Allowances	Denials	
Digestive	3,918	2,322	2,772	1,049	10,061
Diseases of esophagus	17	22	20	13	72
Peptic ulcer (gastric or duodenal)	28	41	21	15	105
Gastritis and duodenitis	(X)	48	44	(X)	128
Hernias	72	160	176	73	481
Crohn's disease	297	266	423	137	1,123
Idiopathic proctocolitis	89	94	114	51	348
Other diseases of gastrointestinal system	397	694	729	291	2,111
Chronic liver disease, cirrhosis	2,968	970	1,224	439	5,601
Gastrointestinal hemorrhage	41	27	(X)	(X)	92
Genitourinary	6,043	500	686	176	7,405
Nephrotic syndrome	219	56	79	23	377
Chronic renal failure	5,731	144	376	36	6,287
Other diseases of the urinary tract	81	175	183	81	520
Disorders of the genital organs	12	125	48	36	221
Skin	255	308	285	122	970
Bullous disease	(X)	(X)	(X)	(X)	13
Ichthyosis	32	56	73	22	183
Dermatitis/psoriasis	80	99	77	26	282
Other disorders of the skin	138	149	133	72	492
Musculoskeletal	46,164	36,793	53,485	21,329	157,771
Diffuse diseases of connective tissue	1,075	483	919	295	2,772
Rheumatoid arthritis	4,138	1,093	1,904	504	7,639
Osteoarthrosis and allied disorders	14,398	6,341	8,852	3,208	32,799
Other and unspecified arthropathies	810	683	705	304	2,502
Ankylosing spondylitis	308	134	222	65	729
Disorders of back (discogenic and degenerative)	19,797	21,150	33,682	13,237	87,866
Disorders of muscle, ligament, and fascia	3,484	5,518	5,696	3,072	17,770
Osteomyelitis and other bone infection	258	86	99	30	473
Other disorders of bone and cartilage	1,761	1,165	1,265	518	4,709
Curvature of spine	135	140	141	96	512
Congenital	123	59	62	33	277
Spina bifida	44	(X)	(X)	(X)	60
Congenital anomalies of heart	60	40	31	19	150
Other congenital anomalies	19	(X)	25	(X)	67

Continued

Table A-1.
Sample distribution, by adjudicative disability category, body system, and primary diagnosis—*Continued*

Body system and primary diagnosis	Initial		Final		Total
	Allowances	Denials not appealed	Allowances	Denials	
Injuries	8,282	8,435	6,632	3,582	26,931
Multiple body dysfunctions	(X)	(X)	(X)	(X)	14
Sleep-related breathing disorders	85	107	150	74	416
Loss of voice	109	21	28	18	176
Fracture of vertebral column	912	108	117	29	1,166
Fracture of upper limb	597	1,069	692	398	2,756
Fracture of lower limb	2,178	2,309	1,836	835	7,158
Other fractures of bones	340	546	425	218	1,529
Dislocations (all types)	104	206	135	63	508
Sprains and strains (all types)	436	2,222	1,407	1,125	5,190
Intracranial injury	593	213	195	72	1,073
Internal injury	10	(X)	17	(X)	41
Open wound, except limbs	(X)	(X)	(X)	(X)	(X)
Open wound upper limb (soft tissue)	216	303	206	117	842
Open wound lower limb (soft tissue)	211	177	146	72	606
Amputations	1,292	770	718	350	3,130
Late effects of injuries to nervous system	1,039	225	301	119	1,684
Chronic fatigue syndrome	102	92	211	67	472
Burns (code 9480)	32	33	26	11	102
Burns (code 9490)	19	22	(X)	(X)	62

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

NOTE: (X) = suppressed to avoid disclosing information about particular individuals.

Table A-2.
Sample distribution, by adjudicative disability category and state

	Initial		Final		
State	Allowances	Denials not appealed	Allowances	Denials	Total
Alabama	3,858	1,583	3,620	625	9,686
Alaska	286	145	85	36	552
Arizona	4,707	1,492	1,725	588	8,512
Arkansas	2,589	913	1,853	533	5,888
California	23,358	10,492	8,279	4,456	46,585
Colorado	2,106	1,368	1,495	507	5,476
Connecticut	2,820	934	1,137	479	5,370
Delaware	828	239	254	115	1,436
Florida	11,372	5,180	8,082	2,839	27,473
Georgia	5,084	2,808	4,310	1,428	13,630
Hawaii	872	286	142	96	1,396
Idaho	918	377	421	174	1,890
Illinois	8,179	3,411	3,794	1,416	16,800
Indiana	4,822	2,555	3,112	1,420	11,909
Iowa	2,339	752	690	386	4,167
Kansas	1,817	814	804	371	3,806
Kentucky	3,552	1,355	3,291	1,119	9,317
Louisiana	2,934	1,388	1,956	709	6,987
Maine	1,313	366	678	183	2,540
Maryland	2,908	1,281	1,500	467	6,156
Massachusetts	5,163	1,280	1,955	646	9,044
Michigan	9,584	4,858	5,087	1,888	21,417
Minnesota	4,209	1,311	1,539	625	7,684
Mississippi	2,343	1,112	1,683	738	5,876
Missouri	5,336	1,846	2,499	846	10,527
Montana	537	291	353	177	1,358
Nebraska	1,261	514	382	221	2,378
Nevada	1,688	529	455	195	2,867
New Hampshire	1,377	320	406	106	2,209
New Jersey	6,863	1,964	2,549	856	12,232
New Mexico	1,285	530	584	238	2,637
New York	15,947	6,143	8,596	2,667	33,353
North Carolina	7,277	3,367	5,064	1,665	17,373
North Dakota	328	150	170	81	729
Ohio	8,028	3,871	4,658	2,150	18,707
Oklahoma	2,834	1,518	2,068	859	7,279
Oregon	2,939	1,278	1,226	644	6,087
Pennsylvania	11,635	4,056	4,866	2,050	22,607
Rhode Island	1,190	282	502	209	2,183
South Carolina	3,769	1,588	2,925	896	9,178
South Dakota	468	199	158	116	941
Tennessee	4,030	1,851	4,182	1,154	11,217
Texas	10,728	5,751	6,669	3,135	26,283
Utah	1,000	486	628	289	2,403
Vermont	490	171	188	72	921
Virginia	5,478	2,096	2,832	1,117	11,523
Washington	4,638	2,048	1,816	786	9,288
West Virginia	2,004	778	2,009	567	5,358
Wisconsin	4,478	1,700	1,664	775	8,617
Wyoming	282	169	171	104	726

SOURCE: Author's calculations based on a 10 percent random sample of the DRF.

Notes

[1] According to the Social Security Advisory Board (2012a), CDRs over the 1996–2008 period resulted on average in more than $10 of savings per $1 spent. Yet, because of budgetary constraints, the number of processed CDRs declined from its peak of more than 1.8 million in 2000 to about 1.1 million by 2009.

[2] In 10 states, a Prototype process initiated in 1999 allows claimants receiving an initial denial to appeal directly to the hearing level without having to go through the reconsideration stage.

[3] The figures in Table 1 are derived from SSA (2009, Tables 60, 61, and 62). Additional years of data appear in those tables. The reason why concurrent applicants are excluded is discussed in the data and methodology section of this article.

[4] The ability to test the impact of any of these factors on the reversal rate of initial denials falls outside the scope of this investigation because of the lack of readily available data. The focus here is on the capacity of primary diagnosis codes to successfully predict disability outcomes through the adjudicative process. A recent preliminary publication by the Social Security Advisory Board (2012b) suggested that third-party representation at the initial determination level increases the likelihood of an allowance substantially for SSI claimants, but only marginally for DI applicants.

[5] For a summary on litigation affecting the disability determination process, see the Social Security Advisory Board (2012a).

[6] Rupp's model did not use the individual primary diagnosis codes, but instead used 16 body systems, which group the specific impairments (15 dummy variables in addition to the musculoskeletal body group serving as the reference category).

[7] Technical denials can occur for a variety of nonmedical reasons, such as engaging in SGA or lacking the required amount of work credits.

[8] For estimation purposes, a 10 percent random sample is used instead of the full DRF because of the computational demands of the estimated models. The 100 percent figures reported in Table 2 are directly derived from the values in Table 1. There are small discrepancies between the two sets of figures. For instance, the 10 percent random sample culls any observations without a known primary diagnosis code or outside the 50 states (Puerto Rico, the District of Columbia, and other territories).

[9] Notice that when estimated from a classical perspective, random coefficient models like the ones in this article make distributional assumptions about subsets of parameters that are in effect no different from those of a prior density. In other words, classical statisticians may also use prior distributions, even if they do not refer to them as such.

[10] All of the models are estimated using Markov Chain Monte Carlo (MCMC) methods. The algorithm is an example of what is known as a Metropolis-within-Gibbs random sampler. A "noninformative" proper prior specification is adopted, with hyperparameter values as suggested by Rossi, Allenby, and McCulloch (2005).

[11] In this article, I focus exclusively on the primary diagnosis codes. A cross-classification of unique primary and secondary diagnosis code combinations would yield many thousands of clusters nesting the individual-level data. Forthcoming research by the author investigates the correlation patterns between primary and secondary diagnosis codes among initial determinations.

[12] To the best of my knowledge, the full extent to which the primary diagnosis change may occur on appeal across the full listing of impairments has never been documented.

[13] Because sex is an individual-level predictor in my models, I merge a few primary impairments that are gender specific. The single category "malignant neoplasm of the genital organs" combines four female diagnosis codes (malignant neoplasms of the uterus, cervix, ovaries, and other female genital organs) with three male diagnosis codes (malignant neoplasms of the prostate, testes, and penis and other male genital organs).

[14] In a Bayesian context, the mean and standard deviation of the posterior density can be used to compute approximate bounds on the posterior probability that a parameter changes sign (much like the t-statistics typically reported in the classical approach).

[15] If a model includes claimant-level predictors, there is a group variance parameter estimate associated with every explanatory variable and not just with the intercepts. However, because the claimant-level predictors have been centered around their grand mean, the intercepts carry the interpretation of adjusted mean linear predictions (see Raudenbush and Bryk (2002)).

[16] In discrete categorical models, a common identification restriction imposes a constant variance. For the multinomial logit case, the within-group variance has a logistic distribution with variance $\pi^2/3$. I follow the approach in Grilli and Rampichini (2007) to recover the ICC estimates.

[17] Notice that a fixed-effects model with the primary impairments rather than body systems would have required 180 indicator variables in the regression, potentially posing serious computational difficulties. In addition, it is unlikely that using the impairments would have substantially increased the share of explained state-level variation.

[18] Surprisingly, as many as 77 percent of the survey respondents were unaware of any activities at the hearing level or above, which appears to undercut the relevance of the result.

References

Coe, Norma B., Kelly Haverstick, Alicia H. Munnell, and Anthony Webb. 2011. "What Explains State Variation in SSDI Application Rates?" CRR Working Paper No. 2011-23. Chestnut Hill, MA: Center for Retirement Research at Boston College.

Congdon, Peter. 2005. *Bayesian Models for Categorical Data*. New York, NY: John Wiley & Sons.

Congressional Budget Office. 2010. "Social Security Disability Insurance: Participation Trends and Their Fiscal Implications." Economic and Budget Issue Brief. Washington, DC: Congressional Budget Office, Health and Human Resources Division (July 22).

Grilli, Leonardo, and Carla Rampichini. 2007. "A Multilevel Multinomial Logit Model for the Analysis of Graduates' Skills." *Statistical Methods & Applications* 16: 381–393.

Hu, Jianting, Kajal Lahiri, Denton R. Vaughan, and Bernard Wixon. 2001. "A Structural Model of Social Security's Disability Determination Process." *Review of Economics and Statistics* 83(2): 348–361.

Keiser, Lael R. 2010. "Understanding Street-Level Bureaucrats' Decision Making: Determining Eligibility in the Social Security Disability Program." *Public Administration Review* 72(2): 247–257.

Lahiri, Kajal, Denton R. Vaughan, and Bernard Wixon. 1995. "Modeling SSA's Sequential Disability Determination Process Using Matched SIPP Data." *Social Security Bulletin* 58(4): 3–42.

Leonesio, Michael V., Denton R. Vaughan, and Bernard Wixon. 2003. "Increasing the Early Retirement Age Under Social Security: Health, Work and Financial Resources." Health and Income Security for an Aging Workforce, Brief No. 7. Washington, DC: National Academy of Social Insurance.

Panis, Constantijn, Ronald Euller, Cynthia Grant, Melissa Bradley, Christin E. Peterson, Randall Hirscher, and Paul Steinberg. 2000. *SSA Program Data User's Manual.* Prepared by the RAND Corporation (contract no. PM-973-SSA) for the Social Security Administration.

Raudenbush, Stephen W., and Anthony S. Bryk. 2002. *Hierarchical Linear Models: Applications and Data Analysis Methods,* 2nd edition. Thousand Oaks, CA: Sage Publications, Inc.

Rossi, Peter E., Greg M. Allenby, and Rob McCulloch. 2005. *Bayesian Statistics and Marketing*. New York, NY: John Wiley & Sons.

Rupp, Kalman. 2012. "Factors Affecting Initial Disability Allowance Rates for the Disability Insurance and Supplemental Security Income Programs: The Role of the Demographic and Diagnostic Composition of Applicants and Local Labor Market Conditions." *Social Security Bulletin* 72(4): 11–35.

Rupp, Kalman, and David Stapleton. 1995. "Determinants of the Growth in the Social Security Administration's Disability Programs: An Overview." *Social Security Bulletin* 58(4): 43–70.

Social Security Administration. 2009. *Annual Statistical Report on the Social Security Disability Insurance Program, 2008*. Washington, DC: Office of Retirement and Disability Policy, Office of Research, Evaluation, and Statistics.

———, Office of the Inspector General. 2010. *Disability Impairments on Cases Most Frequently Denied by Disability Determination Services and Subsequently Allowed by Administrative Law Judges*. Audit Report No. A-07-09-19083. Baltimore, MD: Office of the Inspector General.

Social Security Advisory Board. 2001. *Charting the Future of Social Security's Disability Programs: The Need for Fundamental Change*. Washington, DC: Social Security Advisory Board (January).

———. 2006. *Disability Decision Making: Data and Materials*. Washington, DC: Social Security Advisory Board (May).

———.2012a. *Aspects of Disability Decision Making: Data and Materials.* Washington, DC: Social Security Advisory Board (February).

———. 2012b. *Filing for Social Security Disability Benefits: What Impact Does Professional Representation Have on the Process at the Initial Application Level?* Washington, DC: Social Security Advisory Board (September).

Spiegelhalter, David J., Nicola G. Best, Bradley P. Carlin, and Angelika van der Linde. 2002. "Bayesian Measures of Model Complexity and Fit (with discussion)." *Journal of the Royal Statistical Society (Series B)* 64(4): 583–639.

SSA. *See* Social Security Administration.

Strand, Alexander. 2002. "Social Security Disability Programs: Assessing the Variation in Allowance Rates." ORES Working Paper Series No. 98. Washington, DC: Social Security Administration, Office of Research, Evaluation, and Statistics (August).

The Impact of Retirement Account Distributions on Measures of Family Income

by Howard M. Iams and Patrick J. Purcell*

In recent decades, employers have increasingly replaced defined benefit (DB) pensions with defined contribution (DC) retirement accounts for their employees. DB plans provide annuities, or lifetime benefits paid at regular intervals. The timing and amounts of DC distributions, however, may vary widely. Most surveys that provide data on the family income of the aged either collect no data on nonannuity retirement account distributions, or exclude such distributions from their summary measures of family income. We use Survey of Income and Program Participation (SIPP) data for 2009 to estimate the impact of including retirement account distributions on total family income calculations. We find that about one-fifth of aged families received distributions from retirement accounts in 2009. Measured mean income for those families would be about 15 percent higher and median income would be 18 percent higher if those distributions were included in the SIPP summary measure of family income.

Introduction

The income of the aged is composed largely of three pillars: Social Security benefits, asset income, and pension income (Federal Interagency Forum on Aging-Related Statistics 2012, 14; SSA 2012). In the past three decades, the primary source of pension income has shifted from the traditional defined benefit (DB) pension toward defined contribution (DC) plans, which operate as retirement savings accounts (Anguelov, Iams, and Purcell 2012). The most common DC plans are called 401(k) plans, after the section of the Internal Revenue Code under which Congress first authorized them in 1978.[1] As a consequence of the shift to DC plans, few private-sector employers still offer retirees traditional annuities that provide lifetime income.[2] That trend creates problems for measuring the income of the aged because major government data sources either do not collect information about distributions from retirement accounts or do not include those distributions in their summary measures of income (Anguelov, Iams, and Purcell 2012; Federal Interagency Forum on Aging-Related Statistics 2012, 74).

This article examines the impact of including distributions from retirement accounts on the estimated income of families headed by persons aged 65 or older. After briefly describing our data source, we present our findings in three tables. Table 1 estimates the percentage of families that received distributions from retirement accounts in 2009. Table 2 estimates

Selected Abbreviations

CPS	Current Population Survey
DB	defined benefit
DC	defined contribution
IRA	individual retirement account
IRS	Internal Revenue Service
SIPP	Survey of Income and Program Participation

* *Howard Iams is a senior research advisor to the Office of Research, Evaluation, and Statistics (ORES), Office of Retirement and Disability Policy (ORDP), Social Security Administration (SSA). Patrick Purcell is an economist with the Office of Policy Evaluation and Modeling, ORES, ORDP, SSA.*

the mean and median values of the distributions from retirement accounts. Table 3 estimates the change in family income that would result from including retirement account distributions for affected families. All tables provide breakdowns by age, annual family income (excluding distributions), education, marital status (and, for unmarried persons, sex), and race.[3] We find that about one-fifth of families received distributions from retirement accounts in 2009 and that including those distributions would increase measured mean income for those families by 15 percent and median income by 18 percent. Although the impact of retirement account distributions on retirement income is already significant, it is likely to become even greater in the future as younger cohorts of workers retire after having spent the majority of their careers working at jobs that offered only DC retirement plans.

Data and Methodology

We present data collected in the 2008 panel of the Census Bureau's Survey of Income and Program Participation (SIPP). The data reflect income in 2009, the first full year of income measured for that panel. We focus on the family incomes of married couples, unmarried men, and unmarried women aged 65 or older. SIPP interviews take place every 4 months and collect information about respondents' monthly income in the preceding 4 months. Among other income categories, the SIPP measures the amounts received as distributions from individual retirement accounts (IRAs), Keogh accounts for the self-employed, 401(k)-type DC plans, and lump-sum payments from pension and retirement plans (Census Bureau n.d.).[4] Although the SIPP data file contains amounts received from such distributions each month, its summary measure of total family income excludes those distributions.[5] We summed the monthly values of the retirement plan distributions to estimate the 2009 totals. We then weighted the data using December 2009 weights to represent the US civilian noninstitutionalized population.

We estimate the mean and median values of retirement account distributions for two age groups (65–70, 71 or older) and by quartile of family income (without retirement account distributions), education (high school graduate or less, some college, college graduate), marital status and sex (married, unmarried men, unmarried women[6]), and race (white, black, other). The age categories reflect federal law requiring retirement accountholders to begin taking distributions from IRAs and DC accounts no later than the year after attaining age 70½.[7] The federal required minimum distribution in any year is determined by the account balance and the owner's remaining life expectancy according to Internal Revenue Service (IRS) actuarial assumptions (Purcell 2003). Poterba, Venti, and Wise (2011) analyzed distributions with SIPP data and found that most people did not begin taking distributions from their accounts until they were subject to the required minimum distribution at age 70½.[8]

Results

About 19 percent of families headed by persons aged 65 or older received distributions from retirement accounts in 2009 (Table 1). Retirement distributions were received by a greater share of families headed by persons aged 71 or older (21 percent) than of those aged 65–70 (15 percent). The receipt rate was higher among married couples (25 percent) than among unmarried men (15 percent) and unmarried women (14 percent). Receipt was also more common among families in the fourth (highest) and third income quartiles (24 percent and 25 percent, respectively) than among those in the second and lowest quartiles (18 percent and 8 percent, respectively). Receipt rates increased with educational attainment, ranging from 14 percent among those with a high school education or less to 23 percent among those with some college and to 28 percent among college graduates. Finally, the receipt rate was higher among whites (21 percent) than among blacks (6 percent) or other races (9 percent).

The average value of retirement account distributions received in 2009 by families headed by persons aged 65 or older was $8,121 and the median value was $3,300 (Table 2).[9] The mean value was about two and a half times the median value, suggesting that the amounts were unevenly distributed, with higher values departing much farther from the median than lower values. Average values were higher among families of persons aged 65–70 than those of persons aged 71 or older. The mean distribution amount was higher among married couples ($9,057) than among unmarried men ($7,508) and unmarried women ($6,658). Likewise, the median distribution was higher among married couples ($4,000) than unmarried men ($3,120) and unmarried women ($2,700). Average retirement account distributions increased with family income and education levels. Finally, the mean and median values were higher among families of other races

($11,990 and $4,272, respectively) than were those of whites ($8,116 and $3,400) and blacks ($5,440 and $1,855). The higher values for other races may reflect greater savings rates within that group. [10]

How much would total measured family income increase if distributions from retirement accounts were included? For families who received distributions in 2009, mean family income would increase 15 percent and median income would increase 18 percent if their distributions were included in the SIPP summary measure of total income (Table 3). Mean income would increase from $53,434 without distributions to $61,555 with distributions. Median income would increase by $7,704, from $41,984 without distributions to $49,688 with distributions.

The percent change in mean and median income produced by adding retirement account distributions varies among characteristics and, for some characteristics, the percent change varies between the mean and median values. Mean values are affected by outliers, while a median, representing the middle of the distribution, is unaffected by how extreme the values in the tails of the distribution may be. We found no difference between age groups in the percent change

Table 1.
Families headed by persons aged 65 or older, and percent receiving retirement account distributions, by selected characteristics, 2009

Characteristic	Total families (in thousands)	Families in sample (unweighted)	Percent of families in sample receiving retirement account distributions
Total	24,541	8,080	19
Age			
65–70	8,306	2,747	15
71 or older	16,236	5,333	21
Marital status and sex			
Married couples	10,373	3,425	25
Unmarried men	3,746	1,222	15
Unmarried women	10,422	3,433	14
Age, marital status, and sex			
65–70			
Married couples	4,349	1,427	19
Unmarried men	1,250	410	12
Unmarried women	2,706	910	10
71 or older			
Married couples	6,024	1,998	30
Unmarried men	2,495	812	16
Unmarried women	7,716	2,523	15
Income quartile			
First (lowest)	6,138	2,088	8
Second	6,135	2,064	18
Third	6,134	1,999	25
Fourth (highest)	6,134	1,929	24
Education			
High school or less	14,869	5,044	14
Some college	3,931	1,277	23
College graduate	5,742	1,759	28
Race			
White	21,044	6,819	21
Black	2,308	874	6
Other	1,189	387	9

SOURCE: SIPP, 2008 Panel.

NOTE: Totals do not necessarily equal the sum of rounded components.

of mean family income when including retirement account distributions, but the impact on the median value was higher among the families of persons aged 71 and older (19 percent) than those aged 65–70 (15 percent). The smallest impact on mean and median income by marital status and sex was on married couples (14 percent and 15 percent, respectively) and the largest was on income of unmarried women (18 percent and 20 percent, respectively), with unmarried men falling in between (16 percent and 20 percent, respectively). The effect on mean values was inversely related to family income quartile, falling from 36 percent in the lowest quartile to 26 percent in the second quartile, 18 percent in the third quartile, and 10 percent in the fourth (highest) quartile. The impact on median values also was generally inversely related to income quartile, with the greatest impact on the lowest quartile (18 percent) and the smallest impact on the highest quartile (11 percent). Within educational attainment groups, the greatest impact on mean and median income was among college graduates, although the differences across the education categories were small. Finally, the smallest impact on mean income by race was for black families (7 percent), compared with 16 percent for whites and 15 percent for other races. However, the impact of including retirement account distributions in median family income varied little by race, with all three groups experiencing an increase of 17 to 18 percent.

Table 2.
Families headed by persons aged 65 or older that received retirement account distributions, and mean and median distribution amounts, by selected characteristics, 2009

Characteristic	Families (in thousands)	Families in sample (unweighted)	Mean distribution amount ($)	Median distribution amount ($)
Total	4,620	1,457	8,121	3,300
Age				
65–70	1,231	397	9,720	5,000
71 or older	3,389	1,060	7,541	3,000
Marital status and sex				
Married couples	2,622	847	9,057	4,000
Unmarried men	550	165	7,508	3,120
Unmarried women	1,448	445	6,658	2,700
Age, marital status, and sex				
65–70				
Married couples	806	263	10,580	5,100
Unmarried men	147	45	10,871	5,900
Unmarried women	279	89	6,626	3,075
71 or older				
Married couples	1,816	584	8,382	3,325
Unmarried men	403	120	6,286	2,600
Unmarried women	1,169	356	6,666	2,400
Income quartile				
First (lowest)	519	166	5,283	2,200
Second	1,128	355	6,866	2,800
Third	1,510	477	8,122	3,684
Fourth (highest)	1,463	459	10,095	4,200
Education				
High school or less	2,091	676	6,277	2,400
Some college	922	298	7,026	3,300
College graduate	1,606	483	11,152	4,800
Race				
White	4,363	1,377	8,116	3,400
Black	148	49	5,440	1,855
Other	109	31	11,990	4,272

SOURCE: SIPP, 2008 Panel.

Including distributions from retirement accounts in family income increased mean and median income in all four income quartiles. Among families of persons aged 65 or older, retirement account distributions in 2009 were three times as likely for those in the highest income quartile as for those in the lowest quartile (24 percent versus 8 percent, Table 1). Likewise, retirement account distributions among those in the highest income quartile were substantially larger than those reported in the lowest quartile. In the top quartile, the mean and median total distributions in 2009 were $10,095 and $4,200, respectively, and the corresponding values for the lowest quartile were $5,283 and $2,200 (Table 2).

Although families in the top income quartile were more likely to have received a retirement account distribution, and received larger distributions on average than those in the bottom income quartile, including retirement account distributions in estimates of total income had a negligible impact on income inequality. The share of total income received by people aged 65 or older in the top income quartile fell from 53.9 percent when retirement account distributions were excluded to 53.4 percent when they were included (not shown). The share of total income received by those in the lowest quartile was 7.2 percent, regardless of whether retirement account distributions were included.[11]

Table 3.
Estimated mean and median family income including and excluding retirement account distributions, for families headed by persons aged 65 or older that received distributions, by selected characteristics, 2009

Characteristic	Mean family income: Excluding distributions ($)	Mean family income: Including distributions ($)	Mean family income: Percent increase	Median family income: Excluding distributions ($)	Median family income: Including distributions ($)	Median family income: Percent increase
Total	53,434	61,555	15	41,984	49,688	18
Age						
65–70	63,447	73,166	15	50,994	58,446	15
71 or older	49,796	57,337	15	39,564	46,892	19
Marital status and sex						
Married couples	63,461	72,519	14	52,934	60,840	15
Unmarried men	47,465	54,973	16	35,181	42,171	20
Unmarried women	37,540	44,198	18	27,368	32,952	20
Age, marital status, and sex						
65–70						
Married couples	72,358	82,938	15	57,122	70,623	24
Unmarried men	44,870	55,740	24	36,107	49,035	36
Unmarried women	47,435	54,061	14	31,502	38,805	23
71 or older						
Married couples	59,514	67,896	14	50,676	58,464	15
Unmarried men	48,408	54,694	13	34,651	40,320	16
Unmarried women	35,183	41,849	19	26,681	32,007	20
Income quartile						
First (lowest)	14,619	19,902	36	15,308	18,128	18
Second	26,834	33,700	26	27,034	30,677	13
Third	44,413	52,535	18	43,101	49,712	15
Fourth (highest)	97,023	107,118	10	81,698	90,648	11
Education						
High school or less	43,614	49,891	14	35,814	42,132	18
Some college	48,288	55,314	15	39,386	45,317	15
College graduate	69,177	80,328	16	56,794	67,504	19
Race						
White	52,162	60,278	16	41,652	49,138	18
Black	73,270	78,711	7	49,405	58,405	18
Other	77,545	89,535	15	63,684	74,416	17

SOURCE: SIPP, 2008 Panel.

Are Retirement Account Distributions Income?

The Census Bureau does not measure distributions from retirement accounts in the Current Population Survey (CPS) or the American Community Survey unless they are received as annuities, which are an increasingly uncommon retirement account distribution method (Anguelov, Iams, and Purcell 2012).[12] The SIPP asks about distributions from retirement accounts, but it does not include those distributions in its summary measure of total family income. We believe that, like the SIPP, the CPS and the American Community Survey should collect information about amounts received as distributions from retirement accounts. Then, regardless of whether the Census Bureau includes those distributions in the survey variables that represent total household, family, and personal income, analysts would be able to do so.

Accurately measuring distributions from retirement accounts can be more difficult than measuring income from a DB pension. Typically, DB pension income is received as a monthly annuity. In general, a household survey can ascertain income from a DB pension with three simple questions: Do you receive income from a pension? How often do you receive a pension check? What is the total amount you receive in each check? The same questions can be asked about each DB pension the respondent's household receives.

In contrast to DB pension income, DC account distributions often are taken at irregular intervals, whenever the retiree needs money; or in the case of required minimum distributions, they may occur just once a year. The amount depends on both the account balance and the accountholder's life expectancy, so it changes from year to year. For those reasons, survey respondents may have difficulty recalling distribution amounts and timing. In order to answer those questions accurately, respondents may need to refer to account statements or to the IRS Form 1099-R that they receive each January.

Another complication of counting retirement plan distributions as income is that part of each distribution represents a return to the employee of his or her own prior contributions to the account. In most cases, this problem does not arise with DB pensions because private-sector employees usually do not contribute to their DB plans.[13] Employees' contributions to their retirement accounts were part of their gross income in earlier years, and a general rule of accounting states that a dollar of income in one year should not be counted again as income in a later year.[14] Withdrawals from regular savings accounts, for example, are not treated as income by economists or the IRS because the deposits to those accounts were counted as income in earlier years, as was the interest credited to the account each year. Retirement accounts differ from regular savings accounts in that amounts contributed by employers, and the interest, dividends, and capital gains earned by the account, are not received by the employee until distributions are taken from the account, usually in retirement.

Conclusion

With the shift by employers from providing traditional DB pensions to DC plans over the past several decades, distributions from retirement accounts have become an important resource for the aged. In the private sector, traditional DB pensions that pay lifetime annuities to retirees have been largely supplanted by DC plans, which work like retirement savings accounts. Consequently, a large and growing proportion of Americans are entering retirement with much of their non–Social Security wealth held in retirement accounts. Distributions from those accounts are already a substantial source of income for retirees, and their importance will continue to grow in the future. Consequently, it will be increasingly important for government surveys of household income to accurately measure distributions from those accounts.

We estimate that almost one-fifth (19 percent) of families aged 65 or older received distributions from retirement accounts in 2009.[15] Those distributions had a mean value of $8,121 and a median value of $3,300. If total family income in 2009 as measured in the SIPP had included those distributions, mean income would have been about 15 percent higher and median income would have been about 18 percent higher among families receiving distributions.

As the structure of retirement plans continues to evolve, government surveys that attempt to measure the economic well-being of older persons will need to be revised in response to those changes. If household surveys—especially the CPS, which is used to develop official estimates of household income and the number of persons in poverty—do not accurately identify sources and amounts of income, they will provide misleading results. Inaccurate statistics about household income could lead to inappropriate policies.

Among the Census Bureau's household surveys, the SIPP asks about distributions from retirement plans, but comparisons with IRS data indicate that the SIPP greatly underestimates the amounts of such distributions. The CPS captures distributions from retirement accounts only if they are taken as an annuity, which is not a common form of distribution. Most retirement accountholders take distributions at irregular intervals and in varying amounts. Although distributions from retirement accounts are more difficult to measure than income that is received regularly, the continued relevance of CPS-based estimates of the income of the elderly in the United States depends on the Census Bureau developing appropriate survey questions for that purpose.

Notes

[1] Other employer-sponsored accounts include 403(b) plans for employees of educational and cultural institutions and 457(b) deferred-compensation plans for employees of state and local governments.

[2] In its April–May 2012 survey of employers that sponsor retirement plans, Towers Watson (2012) found that only 6 percent offered a lifetime distribution option, and most of those sponsors reported that less than 5 percent of their employees chose the annuity option at retirement.

[3] We define married couples as those in which the husband is aged 65 or older, and we categorize couples according to the husband's sociodemographic characteristics.

[4] An IRA can contain either a workers' own contributions to the account, amounts that have been "rolled over" into the IRA from a DC plan, or both. The majority of money deposited into IRAs each year consists of rollovers from DC plans (Holden and Schrass 2012).

[5] The SIPP data dictionary defines the income variable TFPNDIST as "family distributions from pension plans: Reaggregated total family distributions from IRA's, KEOGH, and 401(k) pension plans for the reference month after top-coding amounts," and the variable TFLUMPSM as "family retirement lump sum payments: Reaggregated total family lump sum payments from retirement plans for the reference month after top-coding amounts." We sum TFPNDIST and TFLUMPSM to estimate total retirement account distributions. Census Bureau excludes that amount from the variable TFTOTINC, its summary measure of family income.

[6] Unmarried includes never married, widowed, and divorced.

[7] The requirement applies to IRAs and 401(k) plans in which the participant was allowed to defer income taxes on amounts contributed to those plans. Roth IRAs or Roth 401(k) plans require no distributions because contributions to those accounts are taxable in the year they are contributed. In other words, in a traditional IRA or 401(k), income taxes are levied when the money comes out of the account. In a Roth IRA or Roth 401(k), income taxes are levied when the money goes into the account.

[8] Lower-income households with retirement accounts are more likely to take distributions before the required distribution age than are higher-income households. Households in the lower half of the income distribution, however, are less likely to have a retirement account than higher-income households.

[9] Values are calculated only for recipient families; that is, calculations exclude families without retirement account distributions.

[10] Savings tend to rise with income. Asian-Americans constitute the largest group in the "other" race category, and according to the Census Bureau's March 2012 Current Population Survey, the 2011 median household income among Asian-Americans exceeded that of any other race/ethnic group. (DeNavas-Walt, Proctor, and Smith 2012, Table 1).

[11] We had expected that including retirement account distributions in total income would increase income inequality because retirement account ownership is more common in the top income quartile than in the bottom quartile. However, retirement account distributions increased income in almost equal proportions in both quartiles.

[12] Census Bureau officials have indicated that they are considering potential CPS questions about nonannuity retirement account distributions.

[13] With few exceptions, private-sector DB plans are funded by employer contributions and investment earnings. In the public sector, employees usually are required to contribute to their DB pension; therefore, in retirement, some of the income they receive represents a return to them of the contributions they made while they were working. Based on IRS instructions for calculating the taxable portion of pension income received by retirees from public-sector jobs, the return of contributions to retirees usually represents a relatively small fraction of their pension income.

[14] Regardless of whether income taxes are deferred on the employee's contributions, the amount contributed to a DC retirement plan or an IRA is part of his or her gross income in that year.

[15] We believe that 19 percent undercounts the actual share of families receiving such distributions over the year but we do not have access to the data from IRS Form 1099-R, issued by institutions distributing more than $10 from retirement vehicles. The Census Bureau has found that in 2009, about two-thirds of CPS respondents who received 1099-R forms failed to report the distributions in the survey (Bee 2012, Table 2). If that proportion were also to apply to our SIPP data, almost three-fifths of families

would receive distributions, rather than the 19 percent we observe. Bryant, Holden, and Sabelhaus (2011) estimate from tax records that persons older than age 60 in 2006 received about $529 billion in taxable distributions from DC accounts including IRAs. From the SIPP data underlying our calculations for Table 2, we estimate about $144 billion in taxable distributions for families of persons aged 60 or older in 2009, equal to about 27 percent of Bryant, Holden, and Sablehaus' estimate for 2006.

References

Anguelov, Chris E., Howard M. Iams, and Patrick J. Purcell. 2012. "Shifting Income Sources of the Aged." *Social Security Bulletin* 72(3): 59–68.

Bee, C. Adam. 2012. "An Evaluation of Retirement Income in the CPS ASEC Using Form 1099-R Microdata." Paper presented at the Joint Statistical Meetings of the American Statistical Association, San Diego, CA, July 31.

Bryant, Victoria L., Sarah Holden, and John Sabelhaus. 2011. "Qualified Retirement Plans: Analysis of Distribution and Rollover Activity." Pension Research Council Working Paper No. 2011-01. Philadelphia, PA: University of Pennsylvania. http://www.pensionresearchcouncil.org/publications/document.php?file=935.

Census Bureau. n.d. "Survey of Income and Program Participation, 2008 Panel Core File Data Dictionary (All Waves)." http://www.census.gov/sipp/dictionaries/l08puw1d.txt.

DeNavas-Walt, Carmen, Bernadette D. Proctor, and Jessica C. Smith. 2012. *Income, Poverty, and Health Insurance Coverage in the United States: 2011.* Census Bureau, Current Population Reports, P60-243. Washington, DC: Government Printing Office.

Federal Interagency Forum on Aging Related Statistics. 2012. *Older Americans 2012: Key Indicators of Well-Being.* Washington, DC: Government Printing Office. http://www.agingstats.gov/agingstatsdotnet/Main_Site/Data/Data_2012.aspx.

Holden, Sarah, and Daniel Schrass. 2012. "The Role of IRAs in U.S. Households' Saving for Retirement, 2012." *ICI Research Perspective* 18(8). http://www.ici.org/pdf/per18-08.pdf.

Poterba, James M., Steven F. Venti, and David A. Wise. 2011. "The Drawdown of Personal Retirement Assets." NBER Working Paper No. 16675. Cambridge, MA: National Bureau of Economic Research. http://www.nber.org/papers/w16675.

Purcell, Patrick J. 2003. "Retirement Savings Accounts: Early Withdrawals and Required Distributions." *Journal of Pension Planning and Compliance* 29(3): 31–50.

[SSA] Social Security Administration. 2012. *Income of the Population 55 or Older, 2010.* Washington, DC: SSA. http://www.socialsecurity.gov/policy/docs/statcomps/income_pop55/index.html.

Towers Watson. 2012. *Today's Plan for Tomorrow's Retirees: Are We Building DC Plans that Measure Up? 2012 U.S. Defined Contribution Sponsor Survey Report.* New York, NY: Towers Watson. http://www.towerswatson.com/en/Insights/IC-Types/Survey-Research-Results/2012/10/Todays-Plan-for-Tomorrows-Retirees-Are-We-Building-DC-Plans-That-Measure-Up.

Contribution Dynamics in Defined Contribution Pension Plans During the Great Recession of 2007–2009

by Irena Dushi, Howard M. Iams, and Christopher R. Tamborini*

We investigate changes in workers' participation and contributions to defined contribution (DC) plans during the Great Recession of 2007–2009. Using longitudinal information from W-2 tax records matched to a nationally representative sample of respondents from the Survey of Income and Program Participation, we find that the recent economic downturn had a considerable impact on workers' participation and contributions to DC plans. Thirty-nine percent of 2007 participants decreased contributions to DC plans by more than 10 percent during the Great Recession. Our findings highlight the interrelationship between the dynamics in DC contributions and earnings changes. Participants experiencing a decrease in earnings of more than 10 percent were not only more likely to stop contributing by 2009 than those with stable earnings (30 percent versus 9 percent), but they also decreased their contributions substantially (-$1,839 versus -$129). The proportion of workers who decreased or stopped contributions during the crisis exceeded the proportion observed prior to it (2005–2007).

Introduction

Over the past three decades, the pension landscape of the United States has changed dramatically, from one dominated by defined benefit (DB) plans to one where defined contribution (DC) plans are the most prevalent type of retirement plan (Turner and Beller 1989; Gustman and Steinmeier 1992; Employee Benefit Research Institute 1993; Kruse 1995; Rajnes 2002; Costo 2006; Buessing and Soto 2006; Gustman, Steinmeier, and Tabatabai 2009; Purcell 2005, 2009; Copeland 2005, 2009; Bureau of Labor Statistics 2010). This transition has led to a shift of risks and responsibilities from employers to employees who now have to make decisions regarding their own retirement savings. For a DC pension to provide adequate income at retirement, contributions generally need to occur regularly over the work life (Munnell and Sunden 2004). A common view regarding such plans is that once the employee enrolls in the plan and elects his or her contribution amount, inertia will prevail and the employee will continue to contribute in future years.[1]

However, employees may elect to stop, decrease, or increase contributions in any given year in response, among others, to labor market or capital market shocks. Contribution changes that are due to unexpected economic shocks, such as those associated with a recessionary period (for example, housing, income, job and/or financial market shocks), may jeopardize the accumulation of funds in DC retirement accounts and can have an important impact on account balances at retirement, and hence, retirement preparedness. Thus, from a policy perspective it is important to

Selected Abbreviations

DC	defined contribution
SIPP	Survey of Income and Program Participation

** Irena Dushi is an economist with the Office of Policy Evaluation and Modeling, Office of Research, Evaluation, and Statistics (ORES), Office of Retirement and Disability Policy (ORDP), Social Security Administration (SSA). Howard Iams is a senior research adviser to ORES, ORDP, SSA. Christopher Tamborini is a research analyst with the Office of Retirement Policy, ORDP, SSA.*

understand whether and to what extent workers change their contributions over time, particularly in the context of a financial and economic crisis.

This article contributes to the existing literature on the impacts of the economic crisis by investigating the dynamics of employee participation and contributions to DC pension plans during the Great Recession of 2007–2009 and comparing those dynamics with the period prior to it (2005–2007). More specifically, we examine the extent to which changes in contributions are concomitant with earnings changes over the same period.

Using a longitudinal approach, we draw from a data set that links a nationally representative sample of workers from the Survey of Income and Program Participation (SIPP) to their administrative W-2 tax records. These records provide a unique opportunity to examine contribution patterns of the same participants over time. To our knowledge, this study is the first to use a nationally representative sample of individuals matched to administrative records containing longitudinal information about workers' earnings and tax-deferred contributions to examine changes in DC outcomes during the Great Recession.

By examining the impact of the recession on DC pension contributions of the same worker, we provide insights into individuals' responses to economic shocks. Our findings reveal great variability in contributions and indicate that inertia does not typify workers' behavior with respect to contributions to DC plans, especially during the Great Recession. A higher proportion of workers stopped or decreased their contributions substantially (by more than 10 percent) during the recession than did so prior to the recession. Both contribution amounts and contribution rates significantly decreased during the crisis, surpassing in magnitude the slight increase during the period prior to it. Our findings also highlight the role that earnings changes play in altering workers' DC contribution amounts. Thus, workers who experienced decreased earnings were significantly more likely to stop or decrease their contributions than those who did not.

In what follows, we briefly discuss several channels through which the economic downturn may have influenced DC plan contribution behavior and review prior research related to the impact of the Great Recession on DC account activities. Next, we describe our data and empirical strategy and then present our findings from comparing changes in contributions during the crisis with those prior to it. The final section discusses these findings and their implications.

Background

During the 2008–2009 period, the US economy experienced the worst economic downturn since the Great Depression. According to the official definition, the economic downturn, often referred to as the Great Recession, began in December of 2007 and continued through June of 2009 (Business Cycle Dating Committee 2010). The period witnessed rising unemployment, along with falling housing prices, spending, stock prices, household wealth, and retirement assets.

A series of recent studies (Maurer, Mitchell, and Warshawsky 2012; Bricker and others 2011; Butrica, Johnson, and Smith 2012; Johnson and Smith, forthcoming) have revealed substantial impacts of the financial and economic crisis on several outcomes including, spending, retirement plans, and household assets. Hurd and Rohwedder (2010, 2012), for example, found that more than 30 percent of Health and Retirement Study respondents in their fifties decreased their spending during the Great Recession and that the 4–7 percentage point decline in spending was in excess of the decline in previous years. Over 60 percent of families in the Survey of Consumer Finances saw their wealth decline from 2007 to 2009 (Bricker and others 2011). Furthermore, households nearing retirement that were hurt hardest by the dual decline in equity values and home prices changed their retirement behavior in response by increasing saving and deferring retirement (Coronado and Dynan 2012). Given all of these changes, it is plausible that the Great Recession may have also affected participation and contributions to DC pension plans.

Economic and financial downturns may affect workers' retirement savings in employer provided pensions in various ways. Employment and earnings losses, as well as decreasing financial assets, may discourage workers from contributing to a DC pension plan.[2] Furthermore, workers may increasingly prefer to raise their liquid savings outside of retirement accounts during economic downturns, so that savings could be more readily available for consumption if the need arises. At the same time, some workers, particularly those who are not liquidity-constrained, may not change their behavior because of inertia or for other reasons. Others may even increase their contributions because of plan automatic increases or wage increases.

There are several channels through which the Great Recession may have influenced DC pension contribution behavior in the United States. First, a reduction in employment (Hurd and Rohwedder 2010; Coile and Levine 2010) may have put downward pressure on DC participants' contributions. The percentage of the employed population fell from over 63 percent in January 2006 to almost 58 percent by January 2010 (Hall 2010), and the unemployment rate increased from 5 percent in January 2008 to 10 percent by October 2009 (deWolf and Klemmer 2010). Furthermore, labor underutilization increased to 18 percent by the end of 2009, and the number of underemployed workers in part-time jobs rose, mainly reflecting slack demand (Sum and Khatiwada 2010). It is plausible that such employment changes, and the resulting changes in workers' earnings, may have influenced employees' participation and contribution decisions with regard to DC plans.

Second, the financial crisis led to a reduction in employers' matching contributions (Munnell, Aubry and Muldoon 2008a, 2008b). According to the Profit Sharing/401(k) Council of America (2009), during the 2008–2009 downturn, a fifth of private-sector employers either suspended or reduced their matching contributions. In response, employees may have altered their DC contribution amounts.[3] Third, sharp stock market declines and high market volatility may have led to changes in DC contribution behavior. By May of 2009, all retirement accounts had lost $2.7 trillion in assets or 31 percent from their September 2007 peak (Soto 2009).[4] There are other channels, of course, such as changes in household wealth or access to credit during the Great Recession that may have led individuals or households to receive loans or early distributions from their retirement accounts and change their contribution behavior in order to meet debt obligations or consumption needs.[5]

Put together, the economic shocks observed during the Great Recession raise important questions about how employees' contributions to DC plans evolved over the period. To date, despite the critical role that consistency of DC pension contributions plays in retirement security, analyses of DC contribution behavior during periods of labor and financial market shocks are limited, particularly at the population level.[6] A strand of the existing literature uses the administrative records of particular investment firms to analyze cross-sectional aggregates of retirement account activities of account holders during the recession (VanDerhei, Holden, and Alonso 2009, 2010; Holden, Sabelhaus, and Reid 2010). While these studies look extensively at account activities among participants, such as account balances, investment decisions, and participation decisions, they do not link information for the same individual across years and thus do not measure changes in contribution amounts at the individual level. An exception is the recent study by Holden, Sabelhaus, and Reid (2010), which longitudinally tracked account activity of account holders from the beginning of 2008 through September 2009. The authors concluded that only 4.6 percent of plan participants stopped contributions during the first 6 months of 2009, slightly higher than the 3.7 percent of participants in 2008.

Another series of studies by Vanguard—a provider with over 1,100 retirement plans and over a million retirement accounts—also found limited changes in DC participation and contribution rates during the Great Recession (Pagliaro and Utkus 2009a, 2009b; Utkus and Young 2009, 2010; Vanguard 2010). Findings from this set of studies reveal that even though account balances were volatile over the period, the changes in participation and contribution rates among account holders appeared marginal,[7] leading the authors to characterize participants' behavior as driven by inertia (Pagliano and Utkus 2009b).

In sum, prior research using administrative records from retirement investment providers has shown that the majority of participants in DC plans during the Great Recession of 2007–2009 stayed the course and only marginal changes occurred in retirement account activity. However, longitudinal analysis for the same worker over a specified period is limited. It is also unclear from these studies how representative the sample statistics are of all account holders in the United States. Furthermore, the effects of earnings shocks over this period on participation and contributions, while controlling for important demographic covariates and job changes, have not been investigated.

Data and Empirical Strategy

Data for this study come from wave 1 of the 2008 Panel of the Survey of Income and Program Participation (SIPP), which provides us with a nationally representative sample of workers interviewed in the fall, with data collected for the reference period from May through August of 2008, just before the sharp decline in the financial market and job losses associated with the Great Recession toward the end of 2008. While SIPP data provide information about demographic and socioeconomic characteristics of the sample, they do not contain longitudinal information on workers' tax-deferred contributions to retirement accounts. To obtain such data, we match SIPP respondents to their W-2 tax records.[8] These administrative records contain the employer identification number; respondents' annual taxable wage and salary income; and more importantly, tax-deferred contributions to DC accounts over the period of interest in this study (2005–2009). Such information allows us to track job changes, earnings, and tax-deferred contributions to DC plans of the same individuals during the period of Great Recession (2007–2009) and during the immediately preceding period (2005–2007). Another asset of the administrative data, other than their longitudinal feature, is that compared with survey data they provide a more accurate measure of annual earnings and DC pension contributions (Bricker and Engelhardt 2008; Dushi and Honig 2008; Dushi and Iams 2010; Kim and Tamborini 2012).

The analysis sample consists of respondents born from 1949 through 1980 (ages 29–60 in 2009) who according to W-2 records had positive earnings in all 3 years (2005, 2007, and 2009). We select respondents with earnings in those 3 years for two reasons. First, by definition, contributions are tied to employment and earnings; in other words, people with no earnings cannot contribute. Second, we are interested in comparing changes in tax-deferred contributions among wage earners who potentially could have contributed to a plan in both periods: precrisis and during the crisis. While this restriction excludes workers who lost their jobs over each period, our results are not biased because the excluded subsample is comprised of workers with very low earnings, and only a small proportion of them have positive tax-deferred contributions.[9] Another restriction is that respondents must have lived through 2009 to be included in the sample. These restrictions yield an unweighted sample size of 28,128 workers.

Our main goal is to assess whether changes in contributions observed *during the crisis* (2007–2009) exceed those observed during the nonrecessionary period *prior to the crisis* (2005–2007). To do so, we first highlight changes in contributions (both in real dollar amounts and rates) during the crisis and contrast them with similar statistics for the period prior to it.[10] Given our interest in determining the extent to which DC participants changed their contributions because of the recession in excess of what would have been observed in "normal times," we determine the samples for each period separately. Thus, for the period during the crisis, we follow only 2007 contributors through 2009; for the period prior to the crisis, we follow only 2005 contributors through 2007.

Appendix Table A-1 provides characteristics of the entire sample, workers with positive earnings in all 3 years, and separately for those with positive contributions in 2005 (analysis sample for the precrisis period) and in 2007 (analysis sample for the crisis period). Compared with the entire sample of workers, those with positive contributions in 2005 and 2007 are less likely to be female, non-Hispanic blacks, and non-Hispanic others. In addition, contributors are more likely to be married, non-Hispanic whites, and have a college degree or higher level of education.

We first present the distribution of substantial changes in contributions, and their magnitude, over each of the two periods. "Substantial" is considered to be at least a 10 percent change in contributions (in real terms) over the 2-year period, and we classify it into three mutually exclusive categories: decreased by more than 10 percent, increased by more than 10 percent, or stable (within plus/minus 10 percent; that is, contributions remained the same or either decreased by 10 percent or less or increased by 10 percent or less). We measure earnings changes using the same classification as that used for contributions.[11]

Next, we employ multivariate analysis to examine the relationship between the change in DC contributions and earnings changes. We first estimate a *probit* model of the probability of stopping contributions by 2009, where the dependent variable is equal to 1 if the respondent made tax-deferred contributions to an account in 2007 but stopped contributions by 2009, and 0 otherwise.

Then, we estimate Ordinary Least Squares (OLS) regression models of the 2009 tax-deferred contribution amounts and of the 2009 contribution rate. Predictors include a job change variable;[12] log of 2007 earnings; demographic characteristics such as sex, education, marital status, birth cohort, and race/ethnicity, as reported in the 2008 SIPP; and the main variable of interest—the percentage change in earnings from 2007 through 2009. We estimate similar models for the period prior to the crisis, 2005–2007 (available upon request from the authors). Estimates are weighted using SIPP's sampling weights and adjust for its complex sample design.

Finally, we estimate fixed-effect models of the annual DC contribution amounts and of annual contribution rates using person-year panel data from 2005 through 2009. The dependent variable in these models is the contribution amount and, separately, the contribution rate in each year from 2005 through 2009. In these models, we allow DC contributions to be a function of time-varying characteristics such as a job change, real annual earnings, and age at each year. We also allow for time-specific effects by including a dummy variable for each calendar year from 2005 through 2009 that will indicate whether, and to what extent, DC contributions changed in that time period, once we control for the time-varying characteristics. We estimate the OLS models separately for two subsamples: first, we restrict the sample to workers with positive contributions in at least 1 of the 5 years from 2005 through 2009; second, we restrict the sample to workers with positive DC contributions in all of the 5 years over that period.[13] Robust variance estimators are used to correct standard errors for repeated observations of the same individual.

A limitation of the current study, mainly the result of a lack of information in both administrative or survey data, is that it cannot identify the reasons why workers stopped or changed their contributions to DC plans. The observed changes in DC contributions over the period may have occurred for a variety of reasons. They could be involuntary, such as separation from a job or a job loss, a new job that does not offer a DC plan, changes in earnings or employment levels (full or part time), statutory contribution limits, or plan changes such as automatic increases in contributions or changes in the employer match. They could also result from voluntary job changes or be due to a worker's active decision to stop or change contributions. Consequently, although we can estimate the impact of earnings changes on contributions, while controlling for job changes, we cannot tell whether those changes in contributions are due to people making an active or passive decision regarding their savings in tax-deferred plans. Therefore, our findings reveal correlation rather than causality.

DC Contribution Changes During the Great Recession and the Period Prior to It

Table 1 presents the distribution of workers by whether their contributions stopped, remained stable, or substantially increased or decreased during the crisis and contrasts it with the period prior to the crisis. Panel A shows that overall, among 2007 participants, a considerable proportion of them (39 percent) decreased their contributions by more than 10 percent by 2009, including the 16 percent of those who stopped contributing altogether. An additional 32 percent had relatively stable contributions (within plus/minus 10 percent), and the remaining 29 percent increased their contributions by more than 10 percent during the crisis.

As expected, given that contributions are tied to employment and earnings, disaggregating the sample by earnings changes, we observe that for a majority of the sample the change in earnings was accompanied by a similar change in contributions over the same period.[14] Strikingly, 74 percent of workers who saw their earnings decrease by more than 10 percent over the 2007–2009 period had decreased their contributions by more than 10 percent (Table 1, panel A). A significantly larger proportion of 2007 contributors who experienced decreased earnings stopped their contributions by 2009 (30 percent) compared with those with stable earnings (9 percent) or increased earnings (14 percent), suggesting that earnings loss was an important influence.

Panel B presents similar statistics for the period prior to the crisis (2005–2007) and shows considerable fluctuation in contributions even during normal times. Thus, overall, a nontrivial proportion of 2005 contributors (29 percent) decreased their contributions by more than 10 percent by 2007, whereas of the remaining sample about equal proportions had either stable contributions (35 percent) or increased contributions by more than 10 percent (36 percent). Similar to the behavior observed over the 2007–2009 period, 2005 contributors who experienced decreased earnings, compared with those with stable or increased earnings, were significantly more likely to stop or decrease their contributions.

Comparing the two time frames (panel A, the crisis period versus panel B, the precrisis period), reveals that during the crisis, 2007–2009, a statistically significantly higher proportion of workers decreased their contributions by more than 10 percent compared with the period prior to the crisis, 2005–2007 (39 percent versus 29 percent, respectively—a 10 percentage point difference). In addition, a significantly smaller proportion of respondents increased their contributions during the crisis compared with the period prior to it (29 percent versus 36 percent, respectively—a 7 percentage point difference). Furthermore, a significantly higher proportion of workers stopped their contributions during the crisis than in the period before it (16 percent versus 13 percent, respectively). Although the difference between the two periods

Table 1.
Proportion of respondents with positive contributions in the base year, by the magnitude of the change in contributions during and prior to the crisis and earnings changes (in percent)

Earnings change	Decreased by more than 10%	Stable (within plus/minus 10%) [a]	Increased by more than 10%	Total	Stopped contributing by the end of the period	Total N (unweighted)
	Panel A: Crisis period (2007–2009): 2009 contributions relative to those in 2007					
Total	39**	32**	29**	100	16**	12,746
Earnings over the period						
Decreased by more than 10%	74**	14**	12**	100	30	3,286
Stable (within plus/minus 10%) [a]	25**	49	26**	100	9*†	6,006
Increased by more than 10%	28**	20	52**	100	14*†	3,454
	Panel B: Precrisis period (2005–2007): 2007 contributions relative to those in 2005					
Total	29	35	36	100	13	11,560
Earnings over the period						
Decreased by more than 10%	68	17	15	100	30	2,086
Stable (within plus/minus 10%) [a]	19	50	31	100	7†	5,771
Increased by more than 10%	23	21	56	100	12†	3,703

SOURCE: Authors' calculations using Social Security administrative records matched to the 2008 SIPP (wave 1) data.

NOTES: The sample consists of wage and salary workers with positive earnings in all of the 3 years (2005, 2007, and 2009) and with positive contributions in the base year 2007 (or 2005). Reported estimates are weighted.

* denotes that the differences in each cell between the crisis and precrisis periods are statistically significant at the 5 percent level;

** denotes that the differences in each cell between the crisis and precrisis periods are statistically significant at the 1 percent level;

† denotes that the difference within each period between workers who did experience decreased earnings and those with stable (or increased) earnings is statistically significant at the 1 percent level.

a. Contributions (earnings) remained the same or either decreased by 10 percent or less or increased by 10 percent or less.

seems relatively modest (3 percentage points), it represents an increase of 23 percent compared with the precrisis period.

Next, we examine the magnitude of the dollar and percentage change in contribution amounts, as well as in contribution rates during and before the crisis. Note that for each period, we first calculate the change in contributions for each individual and then present the estimated means in Table 2. Panel A shows that during the crisis DC contributions decreased on average by -$399, or by 11 percent.[15] Contributors with decreased earnings of more than 10 percent over the period decreased their contributions substantially, both in real dollars and in percentage terms (on average by -$1,839, or by -46 percent). In contrast, contributors whose earnings increased by more than 10 percent over the crisis period increased their contributions on average by $544, or by 9 percent. With respect to contribution rates, overall they decreased from 6.3 percent of earnings in 2007 to 5.6 percent in 2009, or by 11 percent. The decline in contribution rate was considerable, particularly among workers with decreased earnings (1.4 percentage points, or -26 percent).

In contrast to the crisis period, panel B of Table 2 reveals that overall contribution amounts during the precrisis period increased on average by $121, whereas the contribution rate decreased on average by 3 percent. These changes are significantly different from those observed during the crisis in panel A. During the precrisis period, workers who experienced a substantial decrease in earnings had decreased their contributions on average by -$1,535, or by -39 percent, but these are significantly smaller changes compared with those observed for the similar group during the crisis. In contrast, workers with stable earnings increased their contributions by $263 during the precrisis period compared with a decrease of -$129 during the crisis, leading to a difference-in-difference of -$392. While workers with increased earnings raised their contributions in both periods, the increase was significantly higher during the precrisis period than during the crisis period ($819 versus $544). Charts 1 and 2 depict for

Table 2.
Mean dollar and percentage change of contribution amounts and mean contribution rates and their change during and prior to the crisis among respondents with positive contributions in the base year,[a] by earnings changes

Earnings change	Contribution amount: In the base year (dollars)	Contribution amount: Change over the period [b]: Dollar	Contribution amount: Change over the period [b]: Percent	Contribution rate: In the base year (percent)	Contribution rate: Change over the period [b]: Percentage point difference	Contribution rate: Change over the period [b]: Percent	Total N (unweighted)
			Panel A: Crisis period (2007–2009)				
Total	4,662	-399**	-11**	6.3	-0.7**	-11**	12,746
Earnings over the period							
Decreased by more than 10%	4,745	-1,839*	-46**	5.8	-1.4	-26**	3,286
Stable (within plus/minus 10%) [c]	4,809	-129**†	-2**†	6.7	-0.3**†	-3**†	6,006
Increased by more than 10%	4,321	544**†	9**†	6.1	-0.7**†	-9**†	3,454
			Panel B: Precrisis period (2005–2007)				
Total	4,476	121	0.2	6.2	-0.2	-3	11,560
Earnings over the period							
Decreased by more than 10%	4,493	-1,535	-39	6.0	-1.4	-21	2,086
Stable (within plus/minus 10%) [c]	4,601	263†	4†	6.5	0.2†	3†	5,771
Increased by more than 10%	4,275	819†	16†	6.0	-0.4†	-4†	3,703

SOURCE: Authors' calculations using Social Security administrative records matched to the 2008 SIPP (wave 1) data.

NOTES: The sample consists of wage and salary workers with positive earnings in all of the 3 years (2005, 2007, and 2009) and with positive contributions in the base year 2007 (or 2005). Reported estimates are weighted. Monetary values are in 2009 dollars.

* denotes that the differences in each cell between the crisis and precrisis periods are statistically significant at the 5 percent level;

** denotes that the differences in each cell between the crisis and precrisis periods are statistically significant at the 1 percent level;

† denotes that the difference within each period between workers who did experience decreased earnings and those with stable (or increased) earnings is statistically significant at the 1 percent level.

a. The base year in the crisis period is 2007; in the precrisis period, the base year is 2005.

b. The change in contributions is calculated for each individual, and the reported estimates are the means of the individual changes.

c. Earnings remained the same or either decreased by 10 percent or less or increased by 10 percent or less.

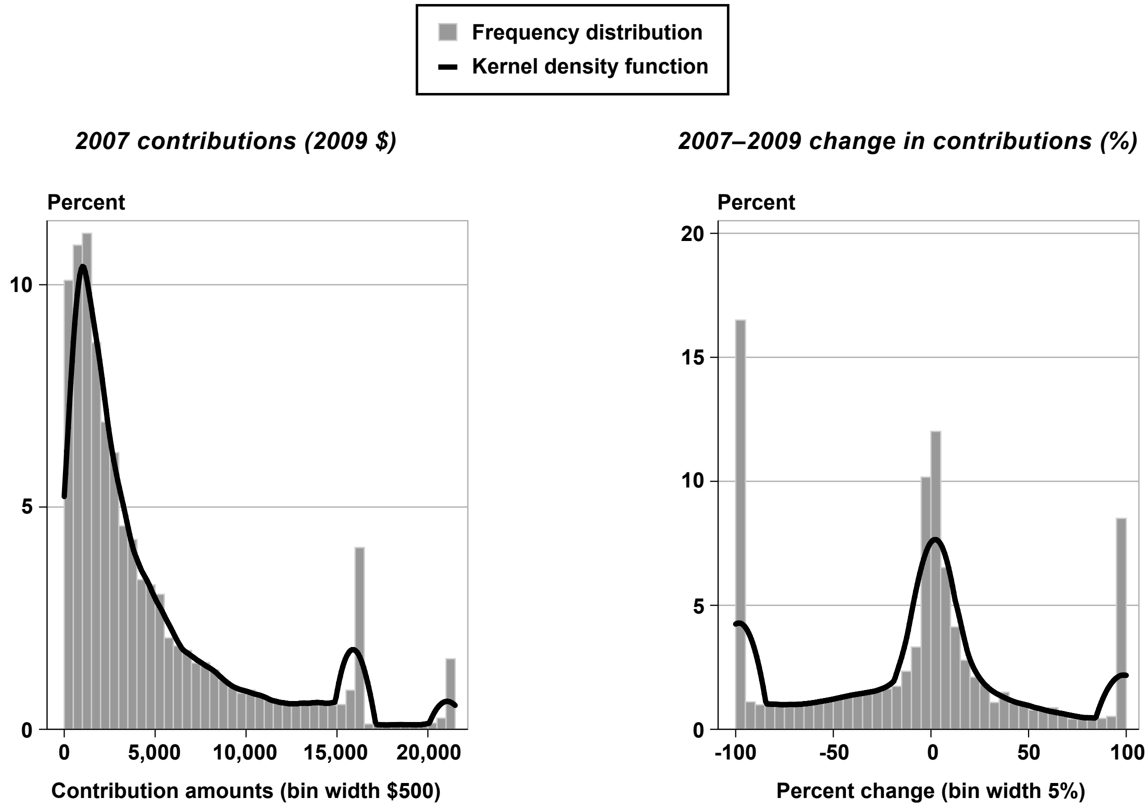
each period (crisis, 2007–2009; precrisis, 2005–2007), respectively, the distribution of contribution amounts in the base year and their percentage change over the period (shown as frequency distributions overlaid by kernel density functions).

Multivariate Estimates of Contribution Changes

We now turn to the multivariate analysis to examine changes in contributions while controlling for observable characteristics. Table 3 (column 1), reports estimated marginal effects of the probability of stopping contributions by 2009.[16] Once we control for observable demographic characteristics and job changes, we observed that workers whose earnings over the period decreased by more than 10 percent were about 71 percentage points more likely to stop their contributions by 2009 than those whose earnings were relatively stable (the omitted category). Workers whose earnings over the period increased by more than 10 percent were about 10 percentage points more likely to stop contributions than those with stable earnings.

Workers with higher 2007 DC contributions had significantly higher contributions in 2009 (Table 3, column 2). Thus, all else equal, a 10 percent higher 2007 contribution leads to an 8 percent higher 2009 contribution. Consistent with the descriptive analysis, respondents who experienced earnings decreases had significantly lower contributions in 2009 (by -$1,534, or 36 percent relative to the mean contribution amount of $4,263), compared with respondents with stable earnings; those who experienced earnings increases had significantly higher contributions in 2009 (by $762, or 18 percent relative

Chart 1.
Distribution of contribution amounts in 2007 and their percentage change during the crisis period (2007–2009)



SOURCE: Authors' calculations using Social Security administrative records.

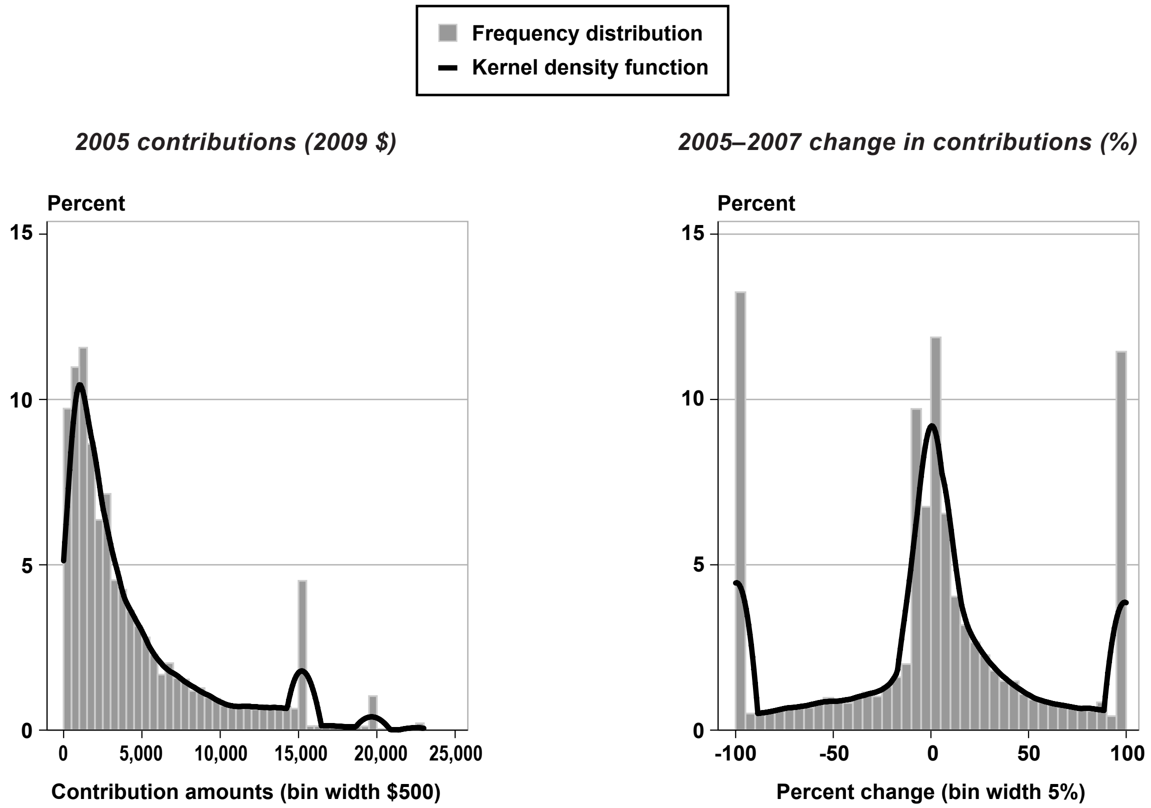
to the mean). Finally, model estimates of contribution rates (column 3), indicate that workers with decreased earnings had significantly lower contribution rates in 2009 (by -.948 percentage points, or 17 percent relative to the mean contribution rate of 5.62) than those with stable earnings; those with increased earnings also had lower contribution rates (by -.235 percentage points, or 4 percent at the mean). It is not surprising to see decreasing contribution rates among workers with earnings gains for two reasons. First, if earnings increased by more than the increase in their contribution amounts, and second, if the majority of those workers have reached the maximum statutory contribution limit, then any wage increases would lead to decreased contribution rates.

Fixed-Effect Models

Overall, the estimated coefficients of the year effects from the fixed-effect models show that annual contributions in real terms increased between 2005 and 2008, but slightly decreased or plateaued in 2009 (see Chart 3 and the Appendix, Table A-2).[17] Thus in 2007, contribution amounts among consistent contributors were significantly higher than in 2005 (by $582, or 11 percent relative to the sample mean of $5,478). In addition, while contributions in 2009 were also significantly higher than in 2005, they were almost the same as those in 2007 or 2008. It is noteworthy that the magnitude of the estimated coefficients is larger among consistent contributors than among those with at least 1 year of contributions, suggesting a greater taste for saving.

Similar patterns of increasing contribution rates between 2005 and 2008 are evident (see Chart 4 and the Appendix, Table A-2). Thus, at the mean, the contribution rate in 2007 among consistent contributors was significantly higher than that in 2005 (0.74 percentage points, or 10 percent relative to the mean contribution rate of 7.05). However, while the contribution rate in 2009 was still significantly higher than that in 2005 (by 0.62 percentage points, or 9 percent relative to the mean), it was significantly lower than that in 2007 (by 0.11 percentage points, or 2 percent). In sum, these findings confirm that contribution patterns

Chart 2.
Distribution of contribution amounts in 2005 and their percentage change during the precrisis period (2005–2007)



SOURCE: Authors' calculations using Social Security administrative records.

during the Great Recession of 2007–2009 differ from the prerecessionary period of 2005–2007. On average, workers increased their contribution rate prior to the recession, but during the recession their contribution rate reversed back to 2007 levels. While at the mean those changes may not seem large, they were greater for a considerable part of the population, as shown in previous tables.

Discussion

Retirement savings in DC pensions represent an increasingly important pillar of retirement security in the United States. This study contributes to the literature by providing insights into the dynamics of workers' contributions to DC plans during the Great Recession of 2007–2009 and comparing those with the period prior to the recession, using longitudinal tax records matched to a nationally representative sample of workers.

Our analysis reveals substantial variability in contributions over multiple years, suggesting that inertia may not typify many workers' DC contribution behavior over time, particularly during the Great Recession. A sizable segment of workers (39 percent) decreased their contributions to DC plans substantially (by more than 10 percent) during the recession. In contrast, during more *normal times*, a significantly lower proportion of workers (29 percent) decreased their contributions substantially. In addition, the proportion of DC participants who stopped contributions during the crisis (16 percent) compared with the period prior to it (13 percent) increased by 23 percent (a 3 percentage point difference). Furthermore, at the mean, both contribution amounts and contribution rates decreased significantly during the crisis of 2007–2009, surpassing in magnitude the increase in contribution amounts and the decline in contribution rates observed during the precrisis period, 2005–2007.

Our findings also highlight the interrelationship between DC contributions and earnings changes. Thus, among workers with positive earnings over the period under study, experiencing a decrease in earnings (whether during or prior to the crisis) has a significant and substantial effect in the likelihood

Table 3.
Probit estimates of the probability of stopping contributions during the crisis period (2007–2009) and OLS estimates of DC plan contributions and contribution rates among respondents with positive contributions in 2007

Independent variable	Probit marginal effects of the probability of stopping contributions by 2009 [a] (1)	OLS regression coefficients: 2009 contribution amount (2)	OLS regression coefficients: 2009 contribution rate [b] (3)
2007 DC plan contributions	-0.00004*	0.799*	---
2007 contribution rate	---	---	0.726*
Log of 2007 annual earnings	-0.228*	515*	0.534*
Earnings change during the crisis period (2007–2009)			
Decreased by more than 10%	0.705*	-1,534*	-0.948*
Stable (within plus/minus 10%) [c]	---	---	---
Increased by more than 10%	0.099*	762*	-0.235*
Constant	1.162*	-4,321*	-3.710*
Predicted mean of dependent variable in 2009	0.119	4,263	5.621
Pseudo R^2 or R^2	0.199	0.707	0.548
N of observations		12,746	

SOURCE: Authors' calculations using Social Security administrative records matched to the 2008 SIPP (wave 1) data.

NOTES: Reported statistics are marginal effects from the probit model and regression coefficients from the OLS model. Control variables include demographic characteristics such as sex, education, birth cohort, race/ethnicity, marital status as reported in the survey year, as well as a dummy variable for at least a job change between 2007 and 2009 generated from the W-2 records. The sample consists of wage and salary workers with positive earnings in all of the 3 years (2005, 2007, and 2009) and with positive contributions in 2007. Standard errors are available from the authors upon request. Reported estimates are weighted and correct for SIPP's complex survey design.

OLS = Ordinary Least Squares;

* denotes statistical significance at the 1 percent level;

--- denotes that the variable is omitted or not included in the regression model.

a. The dependent variable is defined as equal to 1 if the respondent stopped contributing by 2009, and 0 otherwise; the marginal effects are calculated at the sample means and indicate the change in the probability of stopping contributions (in percentage points) for a discrete change in a dummy explanatory variable from 0 to 1, or the change in probability for an infinitesimal change in a continuous explanatory variable.

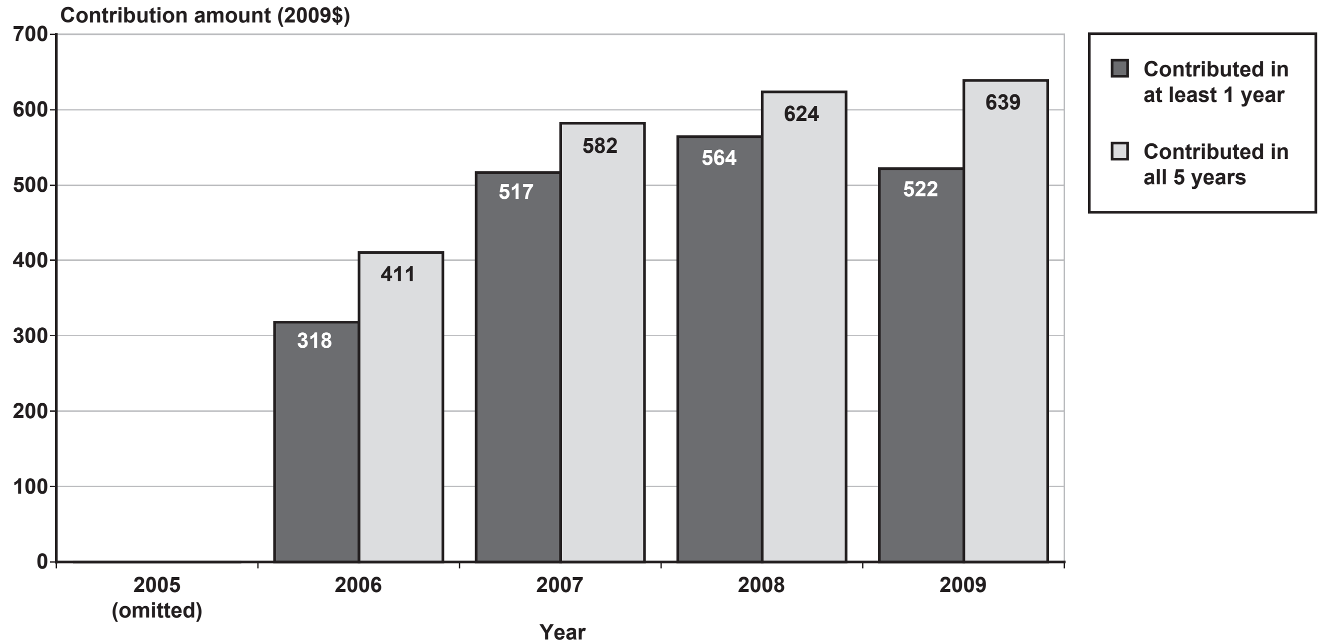
b. The contribution rate is measured as the percentage of annual earnings that are tax-deferred contributions to retirement accounts.

c. Earnings remained the same or either decreased by 10 percent or less or increased by 10 percent or less. This category is omitted.

of stopping contributions by the end of the period. A decrease in earnings also leads to a significant decrease in the contribution amount and contribution rate, suggesting that the loss in earnings is an important factor. Compared to workers with stable earnings, those who experienced an increase in earnings over the period were more likely to stop contributing to their plans. A plausible explanation for this behavior could include unobservable factors such as changes in the employer match, if the respondent is working for a new employer that does not offer a plan, or if the respondent is working for a new employer and is not yet eligible to participate in a plan. In addition, contribution rates declined among workers who experienced earnings increases. A plausible explanation could be that some participants have reached the maximum statutory contribution limit and therefore any wage increases would lead to decreased contribution rates. In sum, these findings suggest that contribution patterns of DC plan participants are quite dynamic and these participants change their contributions (whether voluntary or involuntary) in response to earnings changes.

The findings of this study have important implications for retirement preparedness of employees whose retirement pension income will be drawn mainly from DC pensions. Evidence shows that earnings

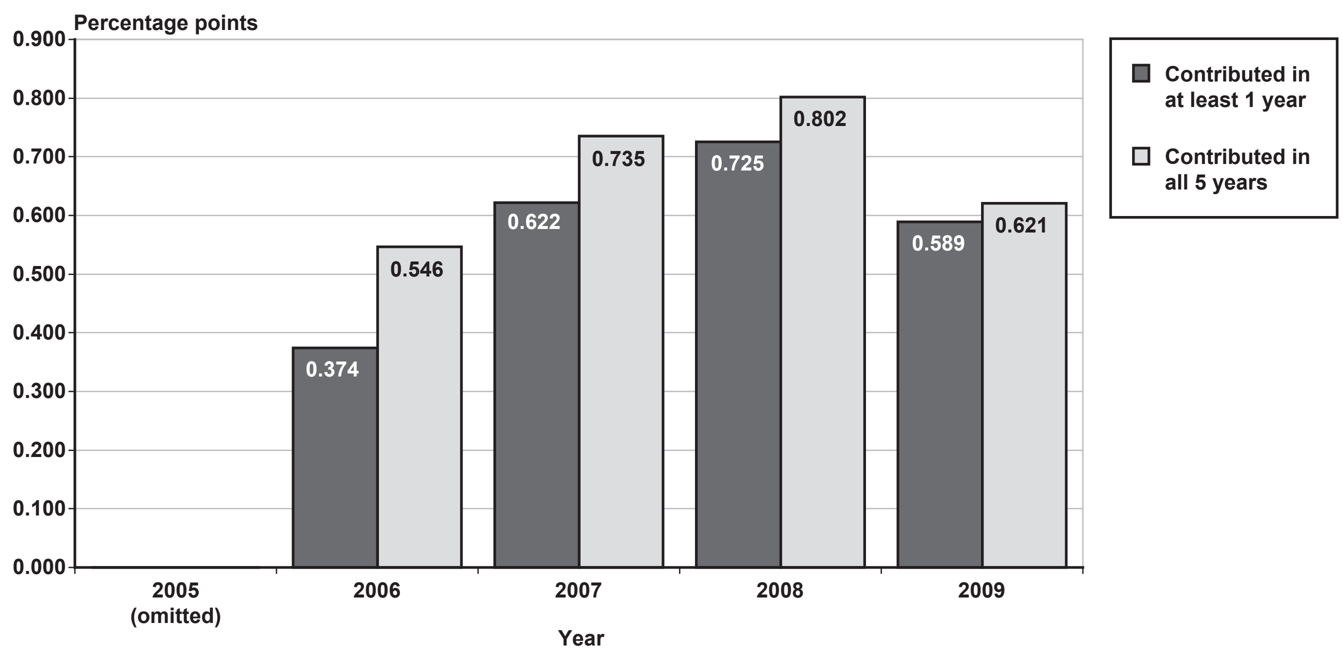
Chart 3.
Coefficient estimates of annual contribution amounts compared with those in 2005, by year



SOURCE: Authors' calculations using Social Security administrative records.

NOTE: All values are statistically significant at the 1 percent level in the given year relative to 2005.

Chart 4.
Coefficient estimates of annual contribution rates compared with those in 2005, by year



SOURCE: Authors' calculations using Social Security administrative records.

NOTE: All values are statistically significant at the 1 percent level in the given year relative to 2005.

shocks that occurred, particularly during the Great Recession, altered workers' participation and contribution amounts to DC plans. Accumulated wealth at retirement will depend not only on the decision to participate in a DC plan and the amount of contributions elected at that time, but will also depend on the employment and earnings shocks experienced throughout one's working life.

Depending on whether the observed changes in contributions are short term or long term, they will have an impact on workers' financial security at retirement. If changes observed over the Great Recession were temporary, then the impact in accumulated assets in DC plans at retirement could be small, whereas a long-term reduction in DC contributions may result in considerably lower retirement wealth. Based on our simulations, assuming that the changes in contributions are temporary, at the mean, account balances at age 62 would be 17 percent lower compared with a "no recession" scenario. However, if those changes were permanent, then their impact could be over 22 percent lower. While it is too early to tell whether the observed changes are temporary or permanent, evidence provided here suggests that researchers should at least be cautious and incorporate such possible changes into their models when making projections of DC pension wealth at retirement.

As noted above, we cannot identify the reasons why workers stopped or changed their contributions to DC plans. The observed changes in DC contributions could be involuntary—such as separation from a job or a job loss, a new job that does not offer a DC plan, changes in employment levels (full or part time), statutory contribution limits, or because of plan changes—such as automatic increases in contributions or changes in the employer match. They could also result from voluntary job changes, or because of a worker's active decision to stop or change contributions. Consequently, although we can estimate the impact on contributions of earnings changes, while controlling for job changes, we cannot tell with certainty whether those changes in contributions are due to people making an active or passive decision regarding their savings in tax-deferred plans. It is plausible that some of those workers may have elected to contribute a percentage of their earnings to their DC plans (about 75 percent of participants according to self-reports in the SIPP data), thus generating automatic increases or decreases in contributions as their earnings changed. If this were the case, then it would suggest that these people did not make an active decision regarding their contributions (that is, a passive change in contributions). However, our results indicate that only about half of workers had a change in contributions of a similar magnitude as that observed in their earnings changes, whereas the remainder of the sample had changes in their contributions in excess of their earnings changes (Table 2). This suggests that they made an active decision.

To further our understanding of whether workers made an active or passive decision regarding their contributions to DC plans, a fruitful avenue of future research may be to examine the effect of a job change on contributions—by comparing workers who change jobs with those who do not change jobs—and its impact on retirement security of different cohorts. Furthermore, it would be valuable to investigate contribution decisions at the household level among married couples because a spouse's contribution decision may respond to the labor market prospect, job changes, pension access, and/or contributions of the other spouse in the household.

Appendix

Table A-1.
Sample characteristics

Characteristic	Sample of all wage earners [a]	Subsample with positive contributions [b]	
		2005	2007
Female	48.5	45.9	46.4
Married	65.7	70.9	70.1
Cohort			
Generation X (born 1965–1980)	48.5	42.7	45.4
Late baby boomers (born 1955–1964)	34.9	38.6	37.4
Early baby boomers (born 1949–1954)	16.6	18.7	17.2
Race/Ethnicity			
Non-Hispanic white	71.5	77.3	76.0
Non-Hispanic black	11.2	9.2	9.6
Hispanic	6.0	6.2	6.3
Non-Hispanic other	11.3	7.4	8.1
Education			
High school graduate or lower	40.4	30.6	31.3
Some college	24.6	24.7	25.0
College graduate or higher	35.0	44.6	43.7
N of observations (unweighted)	28,182	11,560	12,746

SOURCE: Authors' calculations using Social Security administrative records matched to the 2008 SIPP (wave 1) data.

NOTES: Reported estimates are weighted.

a. The sample consists of wage and salary workers with positive earnings in all three years (2005, 2007, and 2009).

b. The subsamples consist of wage earners who contributed to a plan in that year.

Table A-2.
Coefficient estimates from fixed-effect models of the amount of tax-deferred contributions and of contribution rates from 2005 through 2009

Independent variable	OLS model of annual contributions ($)		OLS model of annual contribution rates (%)	
	Sample of contributors [a]	Subsample of consistent contributors [b]	Sample of contributors [a]	Subsample of consistent contributors [b]
Year				
2005	---	---	---	---
2006	318*	411*	.374*	.546*
2007	517*	582*	.622*	.735*
2008	564*	624*	.725*	.802*
2009	522*	639*	.589*	.621*
Overall R^2	0.255	0.457	0.079	0.003
Mean of dependent variable [c]	3,555	5,478	4.85	7.05
Number of person-year observations	79,730	42,200	79,730	42,200
Number of person observations	15,946	8,440	15,946	8,440

SOURCE: Authors' calculations using Social Security administrative records matched to the 2008 SIPP (wave 1) data.

NOTES: The earnings and contributions for each respondent vary by year and are expressed in real 2009 dollars. The estimation controls for other time-varying variables such as age categories, earnings, and job change; it accounts for the fact that there are repeated observations for the same respondent. Robust standard errors are available from the authors upon request. Reported estimates are weighted and account for SIPP's complex survey design.

OLS = Ordinary Least Squares;

--- denotes that the variable is omitted;

* denotes statistical significance at the 1 percent level.

a. The sample consists of wage and salary workers with positive DC contributions in at least 1 of the 5 years from 2005 through 2009.

b. The subsample consists of wage and salary workers with positive DC contributions in all of the 5 years from 2005 through 2009.

c. The mean dependent variable is calculated across all observations in all years.

Notes

Acknowledgments: We thank Paul Davies, Susan Grad, Alan Gustman, Olivia Mitchell, Patrick Purcel, John Sabelhaus, and Alexi Strand for their comments and suggestions to an earlier draft of this article and Karen Smith for her inputs with restricted data.

[1] Findings by Choi and others (2002), for example, suggested that employees often follow the "path of least resistance." Using data from administrative records of several large firms, they showed that the typical employee took over a year to enroll in a 401(k) plan, whereas in companies with automatic enrollment, the majority of employees accepted automatic enrollment defaults such as default saving rates and investment funds.

[2] Chai and others (2012) and Mitchell and Turner (2010) assessed how shocks to human capital shape retirement well-being. The authors showed that human capital risks that are due to fluctuations in labor earnings and unemployment can have profound influence on pension accumulations and thus produce very different pension outcomes.

[3] Munnell and Sunden (2004, 58–60), discussed the impact of employer matching on workers' participation and contribution decisions.

[4] As the stock market recovered, by the first quarter of 2011 retirement account balances were mostly back to their 2007 levels (Butrica and Issa 2011), whereas the unemployment rate and the housing market had not yet recovered.

[5] Note that stock market changes may also lead to changes in contribution behavior. However, we lack information on respondents' asset and portfolio allocation in retirement accounts and their changes over the period, as well as whether observed changes in contributions were in response to stock market shocks.

[6] In a recent paper, Muller and Turner (2011) used longitudinal data from the Panel Study of Income Dynamics to examine the density and persistence of workers' participation in 401(k) plans from 1999 through 2005, but did not look at changes over time in contribution amounts or contribution rates. The authors found that 46 percent of workers who did not change jobs over the period contributed to a plan in all of those years. They concluded that individuals' participation varied over time and that the concept of inertia did not seem to hold for 401(k) saving behavior.

[7] According to their findings, 3.1 percent and 2.9 percent of participants stopped contributions in 2008 and 2009, respectively, compared with approximately 2.5 percent of participants in 2006 and in 2007. In addition, the average contribution rate declined from the 7.3 percent peak in 2007 to 6.8 percent in 2009. In each year from 2006 through 2008, on average, 7 percent of participants decreased their contribution rates.

[8] Olsen and Hudson (2009) and Pattison and Waldron (2008) provide a detail discussion of W-2 tax-record data available in Social Security's Detailed Earnings Records. It is important to note that about 90 percent of adult respondents in the 2008 Panel of SIPP had their survey reports matched to their W-2 records, thus we expect little selectivity bias because of the nonmatch.

[9] From the W-2 records, we can identify a job loss in cases when an individual had positive earnings in a given year but zero earnings in the subsequent year. The W-2 data show that 9.2 percent of all 2007 wage earners lost their jobs by 2009, compared with 6.9 percent of 2005 wage earners who did so by 2007. A very small proportion of contributors, 3 percent and 4 percent (or 330 and 514 observations), respectively, in each period, lost their jobs. Furthermore, in both periods, those who lost their jobs had lower average earnings than those who did not lose their jobs ($12,000 versus $39,000, respectively), suggesting that the excluded group may be comprised of part-time or part-year workers and thus less likely to participate in tax-deferred retirement plans. This analysis (available from the authors on request) indicates that these restrictions do not bias our results and do not considerably understate the decline in contributions; differences in results when including the excluded group in the sample are only trivial.

[10] As noted, the information on contribution amounts is drawn from W-2 records, and thus it is comprised of employee contributions only—the major part of funds invested in DC plans. It is plausible that the magnitude of the change in employee contributions may differ depending on whether or not employers suspended or reduced their matching contributions. However, we have no way of identifying employer contributions or their changes from the administrative or survey data (employer matching contribution is available from the survey at the time of interview, but is not available for the period prior to or after the interview). Broadly speaking, looking at only employee contributions may lead to an overestimate of the decline in contributions among workers whose employer contributions did not change, but to an underestimate among workers whose employer contributions were suspended or reduced.

[11] We selected the 10 percent cut-off point to reflect approximately the average increase in wages over a 2-year period (the annual increase of 5 percent is comprised of both normal wage growth and the inflation rate). In this way, we can distinguish to some extent those changes in contributions that are automatic because of increases in wages and thus may be involuntary (that is, a passive change) from those contribution changes that may be due to substantial wage shocks. Both earnings and contributions are price-indexed to 2009 dollars using the Consumer Price Index for Urban Wage Earners and Clerical Workers (CPI-W) from the *2010 Trustees Report* (Board of Trustees 2010).

[12] Using the employer identification number, we define the job change variable as equal to 1 if in a given year the respondent is working for a new employer, that is, for whom he or she did not work in the previous year. Thus, for the crisis period, the job change dummy variable indicates at least one job change during the 2007–2009 period.

[13] Estimates are reported only for these two samples because we believe they provide the broadest range possible. The first sample allows for workers with earnings to join their plan for the first time or to leave their plan for different reasons (for example, if they changed jobs or became unemployed), and thus it is more representative of the general population. In contrast, the second sample of those with contributions in all 5 years is likely to include longer tenure employees with more stable jobs, and thus it represents a more select sample of workers with DC plans and greater taste for saving.

[14] It is worth noting that if participants elect to contribute to their plan a given percentage of their earnings and do not change it over time, then any increase (or decrease) in earnings will lead to a similar change in contributions without any active decision on their part. Thus, one would expect to see those participants in the diagonal in the table. In contrast, participants with a change in contributions exceeding the change in earnings, suggesting an active decision, would be off the diagonal.

[15] In Table 2, changes in contributions are calculated for each individual, and the reported estimates are the means of the individual changes.

[16] Estimates from the three models for the period prior to the crisis (2005–2007) are similar to those observed during the crisis period (available upon request from the authors).

[17] Please note that samples being analyzed in Appendix Table A-2 and Table 2 differ. In Table 2, we restrict the sample to those with positive contributions in the base year, whereas in Appendix Table A-2, we restrict the sample to consistent contributors (that is, those respondents with positive contributions in all 5 years, columns 2 and 4) and those with contributions in at least 1 year (columns 1 and 3).

References

[Board of Trustees] Board of Trustees of the Federal Old-Age and Survivors Insurance and Federal Disability Insurance Trust Funds. 2010. *The 2010 Annual Report of the Board of Trustees of the Federal Old-Age and Survivors Insurance and Federal Disability Insurance Trust Funds*. Washington, DC: Government Printing Office. http://www.socialsecurity.gov/oact/TR/2010/index.html.

Bricker, Jesse, Brian K. Bucks, Arthur Kennickell, Traci L. Mach, and Kevin Moore. 2011. "Drowning or Weathering the Storm? Changes in Family Finances from 2007 to 2009." NBER Working Paper No. 16985. Cambridge, MA: National Bureau of Economic Research.

Bricker, Jesse, and Gary V. Engelhardt. 2008. "Measurement Error in Earnings Data in the Health and Retirement Study." *Journal of Economic and Social Measurement* 33(1): 39–61.

Buessing, Marric, and Mauricio Soto. 2006. "The State of Private Pensions: Current 5500 Data." Issue in Brief No. 42. Boston, MA: Center for Retirement Research at Boston College.

Bureau of Labor Statistics. 2010. "National Compensation Survey: Employee Benefits in the United States, March 2010." Bulletin No. 2752. http://www.bls.gov/ncs/ebs/benefits/2010/ebbl0046.pdf.

Business Cycle Dating Committee of the National Bureau of Economic Research. 2010. *Report*. Cambridge, MA: National Bureau of Economic Research (September 20). http://www.nber.org/cycles/sept2010.pdf.

Butrica, Barbara A., and Philip Issa. 2011. "Retirement Account Balances." Urban Institute Fact Sheet. http://www.urban.org/publications/411976.html.

Butrica, Barbara A., Richard W. Johnson, and Karen E. Smith. 2012. "Potential Impacts of the Great Recession on Future Retirement Incomes." In *Reshaping Retirement Security: Lessons from the Global Financial Crisis*, edited by Raimond Maurer, Olivia S. Mitchell, and Mark Warshawsky, chapter 3, 36–63. Oxford, UK: Oxford University Press.

Chai, Jingjing, Raimond Maurer, Olivia S. Mitchell, and Ralph Rogalla. 2012. "Life Cycle Impacts of the Financial Crisis on Optimal Consumption—Portfolio Choices, and Labor Supply." In *Reshaping Retirement Security: Lessons from the Global Financial Crisis*, edited by Raimond Maurer, Olivia S. Mitchell, and Mark Warshawsky, chapter 7, 120–150. Oxford, UK: Oxford University Press.

Choi, James J., David Laibson, Brigitte Madrian, and Andrew Metrick. 2002. "Defined Contribution Pensions: Plan Rules, Participation Choices, and the Path of Least Resistance." In *Tax Policy and the Economy,* Vol. 16, edited by James M. Poterba, 67–113. Cambridge, MA: MIT Press.

Coile, Courtney C., and Phillip B. Levine. 2010. *Reconsidering Retirement: How Losses and Layoffs Affect Older Workers*. Washington, DC: Brookings Institution Press.

Copeland, Craig. 2005. "Retirement Plan Participation: Survey of Income and Program Participation (SIPP) Data." EBRI Notes 26(9). Washington, DC: Employee Benefit Research Institute, Education and Research Fund (September).

———. 2009. "Retirement Plan Participation: Survey of Income and Program Participation (SIPP) Data, 2006." EBRI Notes 30(2). Washington, DC: Employee Benefit Research Institute, Education and Research Fund (February).

Coronado Julia L., and Karen Dynan. 2012. “Changing Retirement Behavior in the Wake of the Financial Crisis.” In *Reshaping Retirement Security: Lessons from the Global Financial Crisis*, edited by Raimond Maurer, Olivia S. Mitchell, and Mark Warshawsky, chapter 2, 13–35. Oxford, UK: Oxford University Press.

Costo, Stephanie. 2006. “Trends in Retirement Plan Coverage Over the Last Decade.” *Monthly Labor Review* 129(2): 59–64.

deWolf, Mark, and Katherine Klemmer. 2010. “Job Openings, Hires, and Separations Fall During the Recession.” *Monthly Labor Review* 133(5): 36–44.

Dushi, Irena, and Marjorie Honig. 2008. “How Much Do Respondents in the Health and Retirement Study Know About Their Tax-deferred Contribution Plans?” A Cross-Cohort Comparison.” MRRC Working Paper No. 2008-201. Ann Arbor, MI: University of Michigan Retirement Research Center.

Dushi, Irena, and Howard M. Iams. 2010. “The Impact of Response Error on Participation Rates and Contributions to Defined Contribution Pension Plans.” *Social Security Bulletin* 70(1): 45–60.

Employee Benefit Research Institute. 1993. *Pension Evolution in a Changing Economy*. EBRI Issue Brief No. 141, Special Report SR-18. Washington, DC: Employee Benefit Research Institute (September).

Gustman, Alan L., and Thomas L. Steinmeier. 1992. “The Stampede Toward Defined Contribution Pension Plans: Fact or Fiction?” *Industrial Relations* 31(2): 361–369.

Gustman, Alan L., Thomas L. Steinmeier, and Nahid Tabatabai. 2009. Do Workers Know About Their Pension Plan Type? Comparing Workers’ and Employers’ Pension Information. In *Overcoming the Savings Slump: How to Increase the Effectiveness of Financial Education and Saving Programs*, edited by Annamaria Lusardi, 47–81. Chicago, IL: University of Chicago Press.

Hall, Keith. 2010. “The Economy Today: What Our Measures Tell Us About the Current Labor Market.” Presentation to the Society of Government Economists. Washington, DC: Bureau of Labor Statistics (November 15). http://www.pcbe.org/pcbe.nsf/0/ec2e19ef84b6a078852577d600594a0b/$file/PCBE_KEITH_HALL_OCT_2010.pdf.

Holden, Sarah, John Sabelhaus, and Brian Reid. 2010. *Enduring Confidence in the 401(k) System: Investor Attitudes and Actions*. Washington, DC: Investment Company Institute (January). http://www.ici.org/pdf/ppr_10_ret_saving.pdf.

Hurd, Michael, and Susann Rohwedder. 2010. “Effects of the Economic Crisis and Great Recession on American Households.” NBER Working Paper No. 16407. Cambridge, MA: National Bureau of Economic Research.

———. 2012. “Effects of the Economic Crisis on the Older Population: How Expectations, Consumption, Bequests, and Retirement Responded to Market Shocks.” In *Reshaping Retirement Security: Lessons from the Global Financial Crisis*, edited by Raimond Maurer, Olivia S. Mitchell, and Mark Warshawsky, chapter 4, 64–80. Oxford, UK: Oxford University Press.

Johnson, Richard W., and Karen E. Smith. Forthcoming. “The Great Recession, the Social Safety Net, and Economic Security for Older Americans: Evidence from Multiple National Surveys.” Washington, DC: Urban Institute.

Kim, ChangHwan, and Christopher R. Tamborini. 2012. “Response Error in Earnings: An Analysis of the Survey of Income and Program Participation Matched with Administrative Data.” *Sociological Methods & Research* (November 5). http://smr.sagepub.com/content/early/2012/11/01/0049124112460371.full.pdf+html.

Kruse, Douglas L. 1995. “Pension Substitution in the 1980’s: Why the Shift Toward Defined Contribution?” *Industrial Relations* 34(2): 218–241.

Maurer, Raimond, Olivia S. Mitchell, and Mark Warshawsky. 2012. “Retirement Security and the Financial and Economic Crisis: An Overview.” In *Reshaping Retirement Security: Lessons from the Global Financial Crisis*, edited by Raimond Maurer, Olivia S. Mitchell, and Mark Warshawsky, chapter 1, 1–12. Oxford, UK: Oxford University Press.

Mitchell, Olivia S., and John A. Turner. 2010. “Labor Market Uncertainty and Pension System Performance.” In *Evaluating the Financial Performance of Pension Funds*, edited by Richard Hinz, Heinz P. Rudolph, Pablo Antolin, and Juan Yermo, chapter 5, 119–158. Washington, DC: World Bank.

Muller, Leslie, and John A. Turner. 2011. “The Persistence of Employee 401(k) Contributions Over a Major Stock Market Cycle: Evidence on the Limited Power of Inertia on Savings Behavior.” Upjohn Institute Working Paper No. 11-174. Kalamazoo, MI: Upjohn Institute.

Munnell, Alicia H., Jean-Pierre Aubry, and Dan Muldoon 2008a. “The Financial Crisis and Private Defined Benefit Plans.” CRR Issue Brief No. 8-18. Chestnut Hill, MA: Center for Retirement Research at Boston College.

———. 2008b. “The Financial Crisis and State/Local Defined Benefit Plans.” CRR Issue Brief No. 8-19. Chestnut Hill, MA: Center for Retirement Research at Boston College.

Munnell, Alicia H., and Annika Sunden. 2004. *Coming Up Short: The Challenges of 401(k) Plans*. Washington, DC: Brooking Institute Press.

Olsen, Anya, and Russell Hudson. 2009. “Social Security Administration’s Master Earnings File: Background Information.” *Social Security Bulletin* 69 (3): 29–45.

Pagliaro, Cynthia A., and Stephen P. Utkus. 2009a. "Participant Decisions to Stop Contributions, 2006–2008." Research Note. Valley Forge, PA: Vanguard Center for Retirement Research (February). https://institutional.vanguard.com/iam/pdf/CRRRNC.pdf.

———. 2009b. *Dynamics of Participant Plan Contributions, 2006–2008*. Vanguard Center for Retirement Research, Vol. 37 (August). https://institutional.vanguard.com/iam/pdf/CRRPPC.pdf.

Pattison, David, and Hilary Waldron, 2008. "Trends in Elective Deferrals of Earnings from 1990–2001 in Social Security Administrative Data." Research and Statistics Note No. 2008-03. http://www.socialsecurity.gov/policy/docs/rsnotes/rsn2008-03.html.

Profit Sharing/401(k) Council of America. 2009. *Impact of Economic Conditions on 401(k) and Profit sharing Plans*. Chicago, IL: PSCA. http://www.psca.org/uploads/pdf/research/2009/401k_Economic_Impact_Survey_Final.pdf.

Purcell, Patrick. 2005. *Retirement Plan Participation and Contributions: Trends from 1998 to 2003*. CRS Report for Congress RL33116. Washington, DC: Congressional Research Service (October 12).

———. 2009. *Retirement Plan Participation and Contributions: Trends from 1998 to 2006*. CRS Report for Congress 7-5700. Washington, DC: Congressional Research Service (January 30).

Rajnes, David. 2002. "An Evolving Pension System: Trends in Defined Benefit and Defined Contribution Plans." EBRI Issue Brief No. 249. Washington, DC: Employee Benefit Research Institute (September).

Soto, Mauricio. 2009. "How is the Financial Crisis Affecting Retirement Savings? May 2009, Update." Fact Sheet on Retirement Policy. Washington, DC: Urban Institute. http://www.urban.org/UploadedPDF/901283_retirement_savings_update.pdf.

Sum, Andrew, and Ishwar Khatiwada. 2010. "The Nation's Underemployed in the 'Great Recession' of 2007–09." *Monthly Labor Review* 133(11): 3–15.

Turner, John A., and Daniel J. Beller. 1989. *Trends in Pensions*. Washington, DC: Department of Labor, Pension and Welfare Benefits Administration.

Utkus, Stephen P., and Jean A. Young. 2009. *Inertia and Retirement Savings: Participant Behavior in 2009*. Vanguard Center for Retirement Research, Vol. 36 (April). https://institutional.vanguard.com/iam/pdf/CRRPB.pdf.

———. 2010. *Resilience in Volatile Markets: 401(k) Participant Behavior September 2007–December 2009*. Valley Forge, PA: Vanguard Center for Retirement Research. (March). https://institutional.vanguard.com/iam/pdf/CRRRES.pdf.

VanDerhei, Jack, Sarah Holden, and Luis Alonso. 2009. "401(k) Plan Asset Allocation, Account Balances, and Loan Activity in 2008." EBRI Issue Brief No. 335. Washington, DC: Employee Benefit Research Institute (October).

———. 2010. "401(k) Plan Asset Allocation, Account Balances, and Loan Activity in 2009." EBRI Issue Brief No. 350. Washington, DC: Employee Benefit Research Institute (November).

Vanguard. 2010. *How America Saves 2010: A Report on Vanguard 2009 Defined Contribution Plan Data*. " Vanguard Investment Counseling & Research. Valley Forge, PA: Vanguard Center for Retirement Research. https://institutional.vanguard.com/iam/pdf/HAS.pdf.

OASDI and SSI Snapshot and SSI Monthly Statistics

Each month, the Social Security Administration's Office of Retirement and Disability Policy posts key statistics about various aspects of the Supplemental Security Income (SSI) program at http://www.socialsecurity.gov/policy. The statistics include the number of people who receive benefits, eligibility category, and average monthly payment. This issue presents SSI data for March 2012–March 2013.

The Monthly Statistical Snapshot summarizes information about the Social Security and SSI programs and provides a summary table on the trust funds. Data for March 2013 are given on pages 104–105. Trust fund data for March 2013 are given on page 105. The more detailed SSI tables begin on page 106. Persons wanting detailed monthly OASDI information should visit the Office of the Chief Actuary's website at http://www.socialsecurity.gov/OACT/ProgData/beniesQuery.html.

Monthly Statistical Snapshot

The most current edition of Tables 1–3 will always be available at http://www.socialsecurity.gov/policy/docs/quickfacts/stat_snapshot. The most current data for the trust funds (Table 4) are available at http://www.socialsecurity.gov/OACT/ProgData/funds.html.

Table 1.
Number of people receiving Social Security, Supplemental Security Income (SSI), or both, March 2013 (in thousands)

Type of beneficiary	Total	Social Security only	SSI only	Both Social Security and SSI
All beneficiaries	62,310	54,013	5,511	2,786
Aged 65 or older	40,807	38,718	917	1,172
Disabled, under age 65 [a]	14,112	7,904	4,594	1,615
Other [b]	7,391	7,391	. . .	. . .

SOURCES: Social Security Administration, Master Beneficiary Record and Supplemental Security Record, 100 percent data.

NOTES: Social Security beneficiaries who are entitled to a primary and a secondary benefit (dual entitlement) are counted only once in this table. SSI counts include recipients of federal SSI, federally administered state supplementation, or both.

. . . = not applicable.

a. Includes children receiving SSI on the basis of their own disability.

b. Social Security beneficiaries who are neither aged nor disabled (for example, early retirees, young survivors).

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 2.
Social Security benefits, March 2013

Type of beneficiary	Beneficiaries		Total monthly benefits (millions of dollars)	Average monthly benefit (dollars)
	Number (thousands)	Percent		
Total	57,202	100.0	66,123	1,155.96
Old-Age and Survivors Insurance	46,264	80.9	55,426	1,198.05
Retirement benefits	40,017	70.0	48,797	1,219.43
Retired workers	37,109	64.9	46,973	1,265.82
Spouses of retired workers	2,279	4.0	1,434	629.43
Children of retired workers	629	1.1	390	620.14
Survivor benefits	6,247	10.9	6,629	1,061.10
Children of deceased workers	1,932	3.4	1,548	801.37
Widowed mothers and fathers	146	0.3	130	891.46
Nondisabled widow(er)s	3,912	6.8	4,767	1,218.67
Disabled widow(er)s	256	0.4	182	710.31
Parents of deceased workers	1	(L)	2	1,072.86
Disability Insurance	10,938	19.1	10,697	977.93
Disabled workers	8,852	15.5	10,000	1,129.61
Spouses of disabled workers	160	0.3	48	302.50
Children of disabled workers	1,926	3.4	649	336.81

SOURCE: Social Security Administration, Master Beneficiary Record, 100 percent data.

NOTE: (L) = less than 0.05 percent.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 3.
Supplemental Security Income recipients, March 2013

Age	Recipients: Number (thousands)	Recipients: Percent	Total payments [a] (millions of dollars)	Average monthly payment [b] (dollars)
All recipients	8,298	100.0	4,637	527.51
Under 18	1,312	15.8	865	633.12
18–64	4,897	59.0	2,886	543.95
65 or older	2,089	25.2	886	422.79

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

a. Includes retroactive payments.

b. Excludes retroactive payments.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Trust Fund Data, March 2013

Table 4.
Operations of the Old-Age and Survivors Insurance and Disability Insurance Trust Funds, March 2013 (in millions of dollars)

Component	OASI	DI	Combined OASI and DI
Receipts			
Total	58,016	9,878	67,894
Net contributions [a]	57,098	9,691	66,789
Income from taxation of benefits	15	b	15
Net interest	46	36	81
Payments from the general fund	858	151	1,009
Expenditures			
Total	56,020	12,172	68,192
Benefit payments	55,617	11,833	67,450
Administrative expenses	403	340	743
Transfers to Railroad Retirement	0	0	0
Assets			
At start of month	2,611,311	117,016	2,728,326
Net increase during month	1,996	-2,294	-298
At end of month	2,613,307	114,722	2,728,029

SOURCE: Data on the trust funds were accessed on April 18, 2013, on the Social Security Administration's Office of the Chief Actuary's website: http://www.socialsecurity.gov/OACT/ProgData/funds.html.

NOTE: Totals may not equal the sum of the components because of rounding.

a. Includes reimbursements from the general fund of the Treasury and a small amount of gifts to the trust funds.

b. Between -$500,000 and $500,000.

Supplemental Security Income, March 2012–March 2013

The SSI Monthly Statistics are also available at http://www.socialsecurity.gov/policy/docs/statcomps/ssi_monthly/index.html.

SSI Federally Administered Payments

Awards of SSI Federally Administered Payments

Table 1.
Recipients (by type of payment), total payments, and average monthly payment, March 2012–March 2013

Month	Number of recipients				Total payments [a] (thousands of dollars)	Average monthly payment [b] (dollars)
	Total	Federal payment only	Federal payment and state supplementation	State supplementation only		
2012						
March	8,161,601	5,768,667	2,153,751	239,183	4,507,305	518.60
April	8,185,900	5,980,014	1,981,468	224,418	4,553,734	517.20
May	8,179,285	5,976,689	1,978,456	224,140	4,504,263	516.00
June	8,183,565	5,980,403	1,979,686	223,476	4,494,996	517.80
July	8,225,892	6,014,046	1,988,511	223,335	4,554,428	516.90
August	8,216,619	6,006,681	1,986,567	223,371	4,513,180	517.10
September	8,246,916	6,031,047	1,992,752	223,117	4,515,351	517.70
October	8,277,694	6,055,075	1,999,285	223,334	4,564,279	516.40
November	8,241,018	6,028,214	1,989,793	223,011	4,438,512	518.80
December	8,262,877	6,047,037	1,992,947	222,893	4,593,773	519.43
2013						
January	8,291,772	6,071,217	2,000,021	220,534	4,615,591	525.84
February	8,295,013	6,077,037	1,998,103	219,873	4,612,279	526.41
March	8,297,503	6,079,289	1,998,848	219,366	4,637,309	527.51

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for the end of the specified month.

a. Includes retroactive payments.

b. Excludes retroactive payments.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 2.
Recipients, by eligibility category and age, March 2012–March 2013

Month	Total	Eligibility category		Age		
		Aged	Blind and disabled	Under 18	18–64	65 or older
2012						
March	8,161,601	1,158,789	7,002,812	1,288,548	4,807,814	2,065,239
April	8,185,900	1,156,343	7,029,557	1,301,753	4,821,992	2,062,155
May	8,179,285	1,154,369	7,024,916	1,298,404	4,819,531	2,061,350
June	8,183,565	1,154,725	7,028,840	1,296,051	4,823,143	2,064,371
July	8,225,892	1,157,218	7,068,674	1,305,457	4,849,980	2,070,455
August	8,216,619	1,157,345	7,059,274	1,295,417	4,848,470	2,072,732
September	8,246,916	1,159,205	7,087,711	1,306,587	4,862,627	2,077,702
October	8,277,694	1,161,532	7,116,162	1,309,773	4,884,345	2,083,576
November	8,241,018	1,160,126	7,080,892	1,298,560	4,859,516	2,082,942
December	8,262,877	1,156,188	7,106,689	1,311,861	4,869,484	2,081,532
2013						
January	8,291,772	1,160,197	7,131,575	1,312,233	4,890,028	2,089,511
February	8,295,013	1,157,912	7,137,101	1,316,813	4,890,685	2,087,515
March	8,297,503	1,157,010	7,140,493	1,311,902	4,896,576	2,089,025

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for the end of the specified month.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 3.
Recipients of federal payment only, by eligibility category and age, March 2012–March 2013

Month	Total	Eligibility category		Age		
		Aged	Blind and disabled	Under 18	18–64	65 or older
2012						
March	5,768,667	598,700	5,169,967	1,034,850	3,575,124	1,158,693
April	5,980,014	620,759	5,359,255	1,069,225	3,705,532	1,205,257
May	5,976,689	619,756	5,356,933	1,066,607	3,705,111	1,204,971
June	5,980,403	619,848	5,360,555	1,064,382	3,709,041	1,206,980
July	6,014,046	620,828	5,393,218	1,072,114	3,731,551	1,210,381
August	6,006,681	620,777	5,385,904	1,063,477	3,731,443	1,211,761
September	6,031,047	621,710	5,409,337	1,072,574	3,743,796	1,214,677
October	6,055,075	623,096	5,431,979	1,075,224	3,761,557	1,218,294
November	6,028,214	622,423	5,405,791	1,066,370	3,743,731	1,218,113
December	6,047,037	619,717	5,427,320	1,077,394	3,752,903	1,216,740
2013						
January	6,071,217	622,577	5,448,640	1,077,416	3,770,916	1,222,885
February	6,077,037	621,407	5,455,630	1,081,714	3,773,175	1,222,148
March	6,079,289	620,481	5,458,808	1,077,491	3,779,039	1,222,759

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for the end of the specified month.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 4.
Recipients of federal payment and state supplementation, by eligibility category and age, March 2012–March 2013

Month	Total	Eligibility category		Age		
		Aged	Blind and disabled	Under 18	18–64	65 or older
2012						
March	2,153,751	485,178	1,668,573	252,300	1,110,733	790,718
April	1,981,468	464,224	1,517,244	231,448	1,002,664	747,356
May	1,978,456	463,628	1,514,828	230,607	1,000,704	747,145
June	1,979,686	464,066	1,515,620	230,501	1,000,883	748,302
July	1,988,511	465,637	1,522,874	232,202	1,005,371	750,938
August	1,986,567	465,902	1,520,665	230,737	1,003,971	751,859
September	1,992,752	466,888	1,525,864	232,892	1,006,000	753,860
October	1,999,285	467,938	1,531,347	233,362	1,009,788	756,135
November	1,989,793	467,406	1,522,387	230,977	1,003,014	755,802
December	1,992,947	465,726	1,527,221	233,290	1,004,546	755,111
2013						
January	2,000,021	468,210	1,531,811	233,600	1,007,611	758,810
February	1,998,103	467,285	1,530,818	233,971	1,006,380	757,752
March	1,998,848	467,494	1,531,354	233,335	1,006,735	758,778

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for the end of the specified month.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 5.
Recipients of state supplementation only, by eligibility category and age, March 2012–March 2013

Month	Total	Eligibility category		Age		
		Aged	Blind and disabled	Under 18	18–64	65 or older
2012						
March	239,183	74,911	164,272	1,398	121,957	115,828
April	224,418	71,360	153,058	1,080	113,796	109,542
May	224,140	70,985	153,155	1,190	113,716	109,234
June	223,476	70,811	152,665	1,168	113,219	109,089
July	223,335	70,753	152,582	1,141	113,058	109,136
August	223,371	70,666	152,705	1,203	113,056	109,112
September	223,117	70,607	152,510	1,121	112,831	109,165
October	223,334	70,498	152,836	1,187	113,000	109,147
November	223,011	70,297	152,714	1,213	112,771	109,027
December	222,893	70,745	152,148	1,177	112,035	109,681
2013						
January	220,534	69,410	151,124	1,217	111,501	107,816
February	219,873	69,220	150,653	1,128	111,130	107,615
March	219,366	69,035	150,331	1,076	110,802	107,488

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for the end of the specified month.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 6.
Total payments, by eligibility category, age, and source of payment, March 2012–March 2013 (in thousands of dollars)

Month	Total	Eligibility category		Age		
		Aged	Blind and disabled	Under 18	18–64	65 or older
				All sources		
2012						
March	4,507,305	473,861	4,033,444	840,343	2,805,783	861,179
April	4,553,734	472,480	4,081,255	854,246	2,841,246	858,242
May	4,504,263	471,239	4,033,025	836,006	2,810,846	857,411
June	4,494,996	471,148	4,023,848	840,932	2,795,762	858,301
July	4,554,428	472,715	4,081,712	852,177	2,840,430	861,821
August	4,513,180	472,021	4,041,159	835,979	2,815,453	861,748
September	4,515,351	472,969	4,042,382	843,315	2,808,071	863,966
October	4,564,279	474,596	4,089,683	845,219	2,851,487	867,573
November	4,438,512	472,718	3,965,794	828,040	2,745,321	865,150
December	4,593,773	474,584	4,119,190	856,422	2,867,113	870,238
2013						
January	4,615,591	481,358	4,134,233	856,521	2,875,092	883,978
February	4,612,279	479,815	4,132,464	862,832	2,866,848	882,600
March	4,637,309	481,368	4,155,940	864,978	2,886,289	886,042
				Federal payments		
2012						
March	4,209,479	400,765	3,808,714	826,685	2,640,451	742,343
April	4,269,524	401,949	3,867,575	841,922	2,683,065	744,536
May	4,221,716	400,877	3,820,839	823,837	2,654,041	743,838
June	4,213,739	400,817	3,812,922	828,851	2,640,199	744,689
July	4,270,575	402,084	3,868,490	839,883	2,682,980	747,711
August	4,230,637	401,471	3,829,166	823,909	2,659,044	747,684
September	4,233,203	402,282	3,830,921	831,161	2,652,419	749,624
October	4,279,425	403,684	3,875,742	832,942	2,693,769	752,715
November	4,160,172	402,204	3,757,968	816,241	2,593,035	750,897
December	4,309,786	403,731	3,906,054	844,141	2,710,399	755,246
2013						
January	4,333,173	410,619	3,922,553	844,340	2,719,746	769,087
February	4,331,006	409,172	3,921,834	850,756	2,712,389	767,862
March	4,355,019	410,610	3,944,409	852,896	2,731,132	770,991

(Continued)

Table 6.
Total payments, by eligibility category, age, and source of payment, March 2012–March 2013 (in thousands of dollars)—*Continued*

Month	Total	Eligibility category: Aged	Eligibility category: Blind and disabled	Age: Under 18	Age: 18–64	Age: 65 or older
			State supplementation			
2012						
March	297,826	73,096	224,730	13,658	165,332	118,836
April	284,211	70,531	213,680	12,324	158,181	113,705
May	282,547	70,362	212,185	12,169	156,804	113,574
June	281,258	70,331	210,927	12,082	155,563	113,613
July	283,853	70,631	213,222	12,294	157,450	114,109
August	282,543	70,550	211,993	12,070	156,410	114,063
September	282,148	70,687	211,461	12,154	155,651	114,342
October	284,854	70,912	213,941	12,277	157,718	114,858
November	278,339	70,514	207,826	11,800	152,286	114,253
December	283,988	70,853	213,135	12,281	156,715	114,992
2013						
January	282,418	70,739	211,679	12,181	155,346	114,892
February	281,273	70,643	210,630	12,076	154,459	114,738
March	282,290	70,758	211,532	12,082	155,157	115,050

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for the end of the specified month and include retroactive payments.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 7.
Average monthly payment, by eligibility category, age, and source of payment, March 2012–March 2013 (in dollars)

Month	Total	Eligibility category		Age		
		Aged	Blind and disabled	Under 18	18–64	65 or older
				All sources		
2012						
March	518.60	407.90	536.90	624.90	534.40	415.70
April	517.20	406.90	535.40	621.90	533.00	414.60
May	516.00	407.10	534.00	615.90	532.60	414.70
June	517.80	407.30	535.90	623.70	533.40	414.90
July	516.90	407.20	534.90	619.70	532.80	414.80
August	517.10	407.40	535.20	619.80	533.50	415.00
September	517.70	407.60	535.80	621.30	533.80	415.20
October	516.40	407.50	534.20	614.70	533.30	415.20
November	518.80	407.90	537.00	624.60	534.90	415.60
December	519.43	409.31	537.36	620.77	536.06	416.80
2013						
January	525.84	414.13	544.02	627.01	542.99	422.17
February	526.41	413.41	544.74	631.02	542.93	421.70
March	527.51	414.84	545.78	633.12	543.95	422.79
				Federal payments		
2012						
March	498.40	369.00	519.00	615.70	515.70	379.90
April	498.10	369.10	518.50	613.70	515.20	380.00
May	496.80	369.10	517.00	607.70	514.80	380.10
June	498.60	369.30	519.00	615.60	515.70	380.30
July	497.70	369.10	517.90	611.50	515.10	380.10
August	497.90	369.20	518.20	611.70	515.80	380.30
September	498.50	369.40	518.80	613.20	516.10	380.50
October	497.10	369.20	517.20	606.60	515.50	380.40
November	499.60	369.60	520.10	616.50	517.20	380.80
December	500.29	371.17	520.48	612.68	518.39	382.15
2013						
January	506.75	375.99	527.20	618.83	525.45	387.56
February	507.36	375.16	527.97	622.86	525.43	387.03
March	508.47	376.61	529.02	624.97	526.47	388.15

(Continued)

Table 7.
Average monthly payment, by eligibility category, age, and source of payment, March 2012–March 2013 (in dollars)—*Continued*

Month	Total	Eligibility category: Aged	Eligibility category: Blind and disabled	Age: Under 18	Age: 18–64	Age: 65 or older
			State supplementation			
2012						
March	118.40	129.30	115.10	50.20	124.10	129.80
April	121.90	130.40	119.10	49.00	129.80	131.30
May	121.80	130.40	119.10	49.00	129.70	131.30
June	121.80	130.40	119.10	49.00	129.70	131.30
July	121.70	130.40	119.00	48.90	129.60	131.30
August	121.80	130.30	119.00	48.90	129.60	131.30
September	121.70	130.40	118.90	48.70	129.50	131.30
October	121.70	130.40	118.90	48.70	129.50	131.40
November	121.80	130.40	119.00	48.70	129.60	131.40
December	121.79	130.66	118.95	48.61	129.58	131.56
2013						
January	121.58	130.43	118.75	48.59	129.30	131.38
February	121.47	130.39	118.63	48.48	129.19	131.35
March	121.59	130.51	118.75	48.59	129.27	131.42

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for the end of the specified month and exclude retroactive payments.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

Table 8.
All awards, by eligibility category and age of awardee, March 2012–March 2013

Month	Total	Eligibility category		Age		
		Aged	Blind and disabled	Under 18	18–64	65 or older
2012						
March	79,400	8,823	70,577	15,892	54,531	8,977
April	91,791	9,481	82,310	18,533	63,606	9,652
May	81,195	9,009	72,186	16,222	55,809	9,164
June	76,499	9,105	67,394	15,605	51,675	9,219
July	90,605	9,458	81,147	18,290	62,701	9,614
August	80,464	9,665	70,799	15,810	54,863	9,791
September	77,606	9,462	68,144	14,387	53,623	9,596
October	87,026	9,395	77,631	16,836	60,654	9,536
November	58,337	9,338	48,999	10,868	38,037	9,432
December	82,821	8,679	74,142	16,404	57,626	8,791
2013						
January	72,260	8,293	63,967	14,109	49,729	8,422
February [a]	73,521	9,521	64,000	13,906	49,961	9,654
March [a]	76,196	8,885	67,311	14,349	52,815	9,032

SOURCE: Social Security Administration, Supplemental Security Record, 100 percent data.

NOTE: Data are for all awards made during the specified month.

a. Preliminary data. In the first 2 months after their release, numbers may be adjusted to reflect returned checks.

CONTACT: (410) 965-0090 or statistics@ssa.gov.

PERSPECTIVES—PAPER SUBMISSION GUIDELINES

The *Social Security Bulletin* is the quarterly research journal of the Social Security Administration. It has a diverse readership of policymakers, government officials, academics, graduate and undergraduate students, business people, and other interested parties.

To promote the discussion of research questions and policy issues related to Social Security and the economic well being of the aged, the *Bulletin* welcomes submissions from researchers and analysts outside the agency for publication in its Perspectives section.

We are particularly interested in papers that:

- assess the Social Security retirement, survivors, and disability programs and the economic security of the aged;
- evaluate changing economic, demographic, health, and social factors affecting work/retirement decisions and retirement savings;
- consider the uncertainties that individuals and households face in preparing for and during retirement and the tools available to manage such uncertainties; and
- measure the changing characteristics and economic circumstances of SSI beneficiaries.

Papers should be factual and analytical, not polemical. Technical or mathematical exposition is welcome, if relevant, but findings and conclusions must be written in an accessible, nontechnical style. In addition, the relevance of the paper's conclusions to public policy should be explicitly stated.

Submitting a Paper

Authors should submit papers for consideration via e-mail to Michael V. Leonesio, Perspectives Editor, at perspectives@ssa.gov. To send your paper via regular mail, address it to:

Social Security Bulletin
Perspectives Editor
Social Security Administration
Office of Research, Evaluation, and Statistics
500 E Street, SW, 8th Floor
Washington, DC 20254-0001

We regard the submission of a paper as your implied commitment not to submit it to another publication while it is under consideration by the *Bulletin*. If you have published a related paper elsewhere, please state that in your cover letter.

Disclosures—Authors are expected to disclose in their cover letter any potential conflicts of interest that may arise from their employment, consulting or political activities, financial interests, or other affiliations.

Copyright—Authors are responsible for obtaining written permission to publish any material for which they do not own the copyright.

Formatting Guidelines

To facilitate the editorial process, papers submitted for publication must be prepared in Microsoft Word (**except for tables and charts—see below**) and be formatted as outlined below.

- **Title Page**—Papers must include a title page with the paper's title, name(s) of author(s), affiliation(s), address(es), including the name, postal address, e-mail address, telephone and fax numbers of a contact person. Any Acknowledgments paragraph should also be on this page. In the Acknowledgments, reveal the source of any financial or research support received in connection with the preparation of the paper.
- **Synopsis**—For the *Bulletin's* table of contents include a separate synopsis, including the title of the paper along with one to three sentences outlining the research question.
- **Abstract**—Prepare a brief, nontechnical abstract of the paper of not more than 150 words that states the purpose of the research, methodology, and main findings and conclusions. This abstract will be used in the Bulletin and, if appropriate, be submitted to the Journal of Economic Literature for indexing. Below the abstract supply the JEL classification code and two to six keywords. JEL classification codes can be found at http://www.aeaweb.org/jel/guide/jel.php.
- **Text**—Papers should average 10,000 words, including the text, the notes, and the references (but excluding the tables and charts). Text is double-spaced, except notes and references, which are double spaced only after each entry. **Do not embed tables or charts into the text. Create separate files (in the formats outlined in "Tables/ Charts" below) for the text and statistical material**. Tables should be in one file, with one table per page. Include charts in a separate file, with one chart per page.
- **End Notes**—Number notes consecutively in the text using superscripts. Only use notes for brief substantive comments, not citations. (See the *Chicago Manual of Style* for guidance on the use of citations.) All notes should be grouped together and start on a new page at the end of the paper.
- **References**—Verify each reference carefully; the references must correspond to the citations in the text. The list of references should start on a new page and be listed alphabetically by the last name of the author(s) and then by year, chronologically. Only the first author's name is inverted. List all authors' full names and avoid using *et al*. The name of each author and the title of the citation should be exactly as it appears in the original work.
- **Tables/Charts**—Tables must be prepared in Microsoft Excel. Charts or other graphics must be prepared in or exported to Excel or Adobe Illustrator. The spreadsheet with plotting data must be attached to each chart with the final submission. Make sure all tables and charts are referenced in the text. Give each table and chart a title and number consecutive with the order it is mentioned in the text. Notes for tables and charts are independent of Notes in the rest of the paper and should be ordered using lowercase letters, beginning with the letter a (including the Source note, which should be listed first). The sequence runs from left to right, top to bottom. The order of the notes as they appear below the tables or charts is (1) Source, (2) general notes to the table or chart, if any, and (3) letter notes.

For specific questions on formatting, use the *Chicago Manual of Style* as a guide for notes, citations, references, and table presentation.

Review Process

Papers that appear to be suitable for publication in Perspectives are sent to three reviewers who are subject matter experts. The reviewers assess the paper's technical merits, provide substantive comments, and recommend whether the paper should be published. An editorial review committee appointed and chaired by the Associate Commissioner, Office of Research, Evaluation, and Statistics, makes the final decision on whether the paper is of sufficient quality, importance, and interest to publish, subject to any required revisions that are specified in a letter to the author(s). The entire review process takes approximately 12 weeks.

Data Availability Policy

If your paper is accepted for publication, you will be asked to make your data available to others at a reasonable cost for a period of 3 years (starting 6 months after actual publication). Should you want to request an exception from this requirement, you must notify the Perspectives Editor when you submit your paper. For example, the use of confidential or proprietary data sets could prompt an exemption request. If you do not request an exemption, we will assume that you have accepted this requirement.

Questions

Questions regarding the mechanics of submitting a paper should be sent to our editorial staff via e-mail at ssb@ssa.gov. For other questions regarding submissions, please contact Michael V. Leonesio, Perspectives Editor, at perspectives@ssa.gov.

OASDI and SSI Program Rates and Limits, 2013

Old-Age, Survivors, and Disability Insurance

Tax Rates (percent)	
Social Security (Old-Age, Survivors, and Disability Insurance)	
Employers and Employees, each [a]	6.20
Medicare (Hospital Insurance)	
Employers and Employees, each [a,b]	1.45
Maximum Taxable Earnings (dollars)	
Social Security	113,700
Medicare (Hospital Insurance)	No limit
Earnings Required for Work Credits (dollars)	
One Work Credit (One Quarter of Coverage)	1,160
Maximum of Four Credits a Year	4,640
Earnings Test Annual Exempt Amount (dollars)	
Under Full Retirement Age for Entire Year	15,120
For Months Before Reaching Full Retirement Age in Given Year	40,080
Beginning with Month Reaching Full Retirement Age	No limit
Maximum Monthly Social Security Benefit for Workers Retiring at Full Retirement Age (dollars)	2,533
Full Retirement Age	66
Cost-of-Living Adjustment (percent)	1.7

a. Self-employed persons pay a total of 15.3 percent (12.4 percent for OASDI and 2.9 percent for Medicare).

b. Certain high-income taxpayers will be required to pay an additional Medicare tax beginning in 2013. For details, see the IRS information on this topic (http://www.irs.gov/Businesses/Small-Businesses-&-Self-Employed/Questions-and-Answers-for-the-Additional-Medicare-Tax).

Supplemental Security Income

Monthly Federal Payment Standard (dollars)	
Individual	710
Couple	1,066
Cost-of-Living Adjustment (percent)	1.7
Resource Limits (dollars)	
Individual	2,000
Couple	3,000
Monthly Income Exclusions (dollars)	
Earned Income [a]	65
Unearned Income	20
Substantial Gainful Activity (SGA) Level for the Nonblind Disabled (dollars)	1,040

a. The earned income exclusion consists of the first $65 of monthly earnings, plus one-half of remaining earnings.

Social Security Administration
Office of Retirement and Disability Policy
Office of Research, Evaluation, and Statistics
500 E Street, SW, 8th Floor
Washington, DC 20254

SSA Publication No. 13-11700
May 2013