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# Social Security Bulletin

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*by Patrick J. Purcell*

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# Geographic Mobility and Annual Earnings in the United States

by Patrick J. Purcell*

*The geographic mobility rate of U.S. workers has declined in recent decades. Labor mobility has historically indicated variations between local areas in earnings and other economic conditions. Because average career earnings determine Social Security retirement benefit levels, changing trends in geographic mobility and earnings may have implications for workers' future benefits. I use administrative data on earnings from the Social Security Administration's Continuous Work History Sample to examine trends in geographic mobility from 1994 to 2016 and to compare the earnings of working-age adults who moved to another county or state with the earnings of those who did not. I find that the relative difference in earnings between movers and nonmovers changed little during the observation period. Although some researchers have suggested that declining labor mobility has resulted from a decline in the earnings gains workers can realize by moving, this finding suggests that such a link is unlikely.*

## Introduction

The economy of the United States is highly dynamic in terms of the number of jobs it creates and the ease with which workers move from one job to another (Hyatt and others 2018). A key element of this dynamism is the willingness of workers to relocate in response to geographic differences in wages and employment rates. Consequently, Americans are often perceived as being more geographically mobile than residents of other developed nations (Pingle 2007; Frey 2009). This perception is supported by evidence that historical internal migration rates have been higher in the United States than in most other developed countries (Cooke 2011; Molloy, Smith, and Wozniak 2011; Partridge and others 2012). Worker mobility plays an important role in mitigating geographic differences in employment and earnings (Pingle 2007; Levy, Mouw, and Perez 2017). For example, when workers respond to adverse income shocks by relocating, they also promote a macroeconomic adjustment to regional downturns in employment (Bayer and Juessen 2012). Through this mechanism, geographic mobility "has been shown to smooth out spatially-asymmetric macroeconomic shocks and the effects of industry restructuring" (Partridge and others 2012).

Researchers have noted a decline in the geographic mobility of Americans of working age over the last several decades and have suggested several explanations for this trend, including the possibility that the earnings gains movers can realize have declined over time. Some researchers have also found evidence of a coinciding decline in the rate at which workers change employers (with or without relocating). To date, the relationship between geographic mobility, employer change, and earnings remains largely unexplored because the data underlying such research have typically come from household surveys, few of which follow a large sample of individuals over a long period. However, an administrative data file such as the

**Selected Abbreviations**

| | |
|---|---|
| CWHS | Continuous Work History Sample |
| OLS | ordinary least square |
| SSA | Social Security Administration |

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Social Security Administration's (SSA's) Continuous Work History Sample (CWHS) includes long-term information on workers' earnings, their employers, and their county of residence. Thus, the CWHS is ideally suited for a study of trends in geographic mobility, employer change, and earnings.

Policymakers might ask whether the decline in geographic mobility is cause for concern. Because labor mobility has been a means through which workers have adjusted to adverse shocks in local employment and earnings (Blanchard and Katz 1992), declining mobility rates could lengthen spells of unemployment if unemployed workers are less likely to move to places with higher labor demand. On the other hand, geographic mobility may have declined because improvements in information and communications technology have enabled better matching of individuals to jobs within local labor markets (Molloy, Smith, and Wozniak 2017).

This article uses data from the CWHS for the period from 1994 through 2016 to address the following questions:

- What do administrative earnings records from SSA reveal about trends in geographic mobility and changing from one employer to another over the observation period?
- What personal characteristics and local economic variables are associated with mobility and employer change, and did the statistical relationships change during this period?
- How do the earnings of movers and nonmovers compare, and did the relative earnings of movers and nonmovers change during this period?

This information may help policymakers better understand how geographic mobility is related to workers' earnings, which ultimately determine Social Security benefit levels.

This article is organized into eight sections, including this introduction, arranged as follows:

- The three sections that follow this introduction review the relevant literature, describe the data and methods used in this study, and present descriptive statistics on geographic mobility and employer change in three 3-year periods (1994–1996, 2004–2006, and 2014–2016).
- The fifth section presents the results of logistic regressions examining the statistical relationships between mobility and selected personal and geographic variables; and between employer change and those variables.
- The sixth section presents statistics on the average annual earnings of individuals before and after relocating and compares them with the earnings of workers who, in the same period, did not relocate.
- The seventh section presents the results of a regression analysis examining the statistical relationship between earnings change over time and selected personal and geographic variables.
- The eighth section summarizes and concludes.

## Previous Studies

Geographic mobility in the United States has declined steadily for more than 40 years (Frey 2009; Cooke 2011; Partridge and others 2012; Molloy, Smith, and Wozniak 2017). Analysts have identified several possible causes for the long-term decline in migration. Pingle (2007) suggested that lower levels of migration in the 1980s and 1990s were due in part to a reduction in military enrollment, given that servicemembers move more often than civilians do. Frey (2009) attributed much of the decline to increases in the median age of the population and in homeownership rates. Winkler (2011) found that homeownership makes workers less likely to move in response to labor market shocks. Cooke (2013) suggested that an increasing prevalence of two-earner couples, greater household indebtedness, and the development of modern information and communication technologies have contributed to lower rates of geographic mobility. Karahan and Rhee (2014) estimated that the increasing median age of the population, and that trend's secondary effects on the labor market, could account for about half of the long-term decline in geographic mobility in the United States.[1] Foster (2017) attributed up to one-third of the long-term decline in geographic mobility to increases in the median age and the share of the population comprising historically less-mobile racial/ethnic groups, but concluded that rising homeownership rates and the increasing prevalence of dual-earner couples had negligible effects on mobility.

In contrast with the studies cited above, Molloy, Smith, and Wozniak (2017) concluded that older median ages, rising homeownership rates, and other observable demographic and socioeconomic factors played minimal roles in the decline in mobility. Kaplan and Schulhofer-Wohl (2017) found that neither older median ages nor the greater prevalence of two-earner households could explain the decline in mobility. They estimated that convergence in regional wage rates and improvements in information and

communications technology might explain more than half of the decline in interstate migration rates from 1991 to 2011.

Some researchers have identified a downward trend in the movement of workers from one employer to another which coincides with the decline in geographic mobility, and several have suggested that the two declines are related. Examining data covering the late 1960s to the late 2000s, Molloy, Smith, and Wozniak (2014) suggested that the most plausible reason for both declines was a reduction in the earnings gains from making such transitions and that the decline in financial gains from changing employers caused the decline in geographic mobility. Then, in their 2017 article, the authors again contended that declining rates of employer change and geographic mobility were related and that the former caused the latter. Examining data from the Panel Study of Income Dynamics, Gittleman (2019), too, observed a decline in employer-change transitions and estimated that the increasing median age of the workforce was responsible for about three-fifths of the decline. By contrast, Hyatt and others (2018) concluded that declines in earnings gains from migration explained little of the long-term decline in geographic mobility.

The studies cited above demonstrate the disagreement about the causes of declining geographic mobility in the United States over the last several decades. Although some researchers have attributed the decline mainly to a rising median age and increasing homeownership rates, others have suggested that convergence in regional wage rates and improvements in communication technology are more likely causes. Furthermore, not all research has found a decline in employer-change transitions concurrent with the more widely documented decline in geographic mobility.

People who move to a new geographic area are not a random cross-section of the U.S. population (Borjas, Bronars, and Trejo 1992). Several studies have found evidence of self-selection among movers based on age, education, and annual earnings. Gabriel and Schmitz (1995) and Rodgers and Rodgers (2000) found that workers who moved to a new location had higher annual earnings prior to moving relative to nonmovers with similar characteristics. Dahl (2002), Wozniak (2010), and Levy, Mouw, and Perez (2017) all found that college-educated individuals are more geographically mobile than those who did not attend college. Kennan and Walker (2011) reported that younger and more educated people are more likely to move to a new area than those who are older and less educated. They also noted that multiple movers and returning movers account for a large share of moves. Coen-Pirani (2010) observed that recent immigrants to the United States migrate to new locations more frequently than nonimmigrants do, and that differences in geographic mobility rates across states are not fully explained by differences in age and education. At least one study (Yankow 2004) found that unemployed persons are significantly more likely to move than employed workers are.

Just as movers are not randomly selected from the population, neither are the locations to which they relocate. In general, people tend to move to locations that pay higher wages for their particular skills (Borjas, Bronars, and Trejo 1992). Kennan and Walker (2011) observed that geographic differences in average wages are a significant determinant of where workers choose to live. Likewise, Kaplan and Schulhofer-Wohl (2017) concluded that movers tend to go to states where their particular occupations are better paid. The systemic differences between movers and nonmovers, and between places that attract movers and those that do not, have earnings implications. Dahl (2002) estimated that self-selection of more educated workers to states with higher returns to education can bias the estimated return on a college education upward by 10–20 percent. In addition to being more likely to move than are those with less education, college graduates respond more to differences in local labor market conditions when choosing where to live (Wozniak 2010). Levy, Mouw, and Perez (2017) also found that wage- and unemployment-rate differences substantially affect the destination choices of workers who move. Ganong and Shoag (2017) found that rising housing prices in high-income areas, by eroding the gains from moving, have deterred moves among low-skill workers.

Both the self-selection of movers and the higher average wages in their chosen destinations can affect postmove earnings; however, because of limitations in the available data sets, relatively few studies have examined the postrelocation earnings of movers. Using data from the Current Population Survey (CPS) Displaced Workers Files, Raphael and Riker (1999) concluded that geographic mobility has a substantial and significant positive effect on the earnings of displaced workers.[2] Yankow (1999) studied data from the National Longitudinal Survey of Youth (NLSY) and found that young interstate migrants realized significant earnings gains over the 5-year period following

a move and concluded that the earnings gain was not caused by movers being disproportionately drawn from the upper tail of the distribution of skills and abilities. Rodgers and Rodgers (2000) analyzed data from the Panel Study of Income Dynamics and estimated that men's real earnings 6 years after moving were 20 percent higher than they would have been otherwise. The authors also found that almost all earnings gains occurred among men who were younger than 40 in the year they moved. Kennan and Walker (2011) studied data from the NLSY and found that among white men with a high-school education, expected gains in earnings influence geographic mobility, but the analysis did not include measures of postrelocation earnings. Bayer and Juessen (2012) combined data from the CPS with administrative data from the Internal Revenue Service and found persistent income gains from geographic mobility.

## Data and Methods

The present analysis was conducted on individual earnings histories from the CWHS, which contains earnings records that represent 1 percent of all Social Security numbers (SSNs) ever issued. To maintain the CWHS at 1 percent of all SSNs, SSA adds the earnings records of a random selection of newly issued SSNs each year. The records of deceased workers remain in the CWHS, allowing researchers to study the annual wages of entire birth cohorts over time. When necessary, SSA updates the CWHS earnings records for adjustments and corrections to SSA's Master Earnings File. For research purposes, the CWHS—with its large number of earnings records, longitudinal structure, and accuracy—has several advantages over household surveys. Specifically, most household surveys consist of smaller samples, collect data for relatively short periods, and are subject to participant nonresponse and recall errors.

The CWHS includes data on Social Security taxable wages in covered employment from 1951 forward. Covered employment refers to jobs (or self-employment) subject to Social Security payroll-tax deductions. Wages in covered employment are taxable up to an annually adjusted threshold amount called the taxable maximum. Workers' taxable wages in covered employment are the basis on which SSA determines both eligibility for Social Security benefits and the amounts of those benefits. Prior to 1978, the CWHS tracked only covered earnings; since then, it has also included annual wages in noncovered employment and earnings above the annual maximum taxable amount.[3]

This article describes results derived from the 2016 CWHS, the most recent file available when the analysis was conducted. The 2016 CWHS includes 3,467,451 individual person-records, of which 52.0 percent are for men and 48.0 percent are for women. The earnings analyzed in this article consist of annual wages and salaries in both covered and noncovered employment, including those exceeding the annual taxable maximum. Self-employment earnings are also included. Men's and women's annual earnings are analyzed separately. I restrict the analysis to earnings accrued from ages 25 through 49, which are the ages with the highest employment rates and the highest rates of geographic mobility (Bureau of Labor Statistics 2019). For brevity, I refer hereafter to all wage, salary, and self-employment income simply as "earnings." All annual earnings have been indexed to 2016 values using the Consumer Price Index for All Urban Consumers.

Geographic mobility can be defined several ways. The three most common definitions identify a move as relocating either to a different state, to a different county *or* state, or to a different commuting zone.[4] In this article, I define geographic mobility as moving to a different county or state. Individuals who move to a new address in the same county are not considered movers. One benefit of defining geographic mobility at the county level is that the 3,142 counties and county-equivalents in the United States range from the rural and sparsely populated to the urban and densely populated. Moreover, counties (unlike commuting zones) have stable borders that are not affected by population growth (Partridge and others 2012). This is helpful for studying long-term trends. Finally, many of the local-area economic statistics that indicate the factors that may influence a worker's decision to move to a new location—including median household income and unemployment rate—are available at the county level. Thus, in this article, the terms "movers," "mobility," and "relocation" refer to workers who moved to another county or state.

Another aspect of mobility is its duration. To compare the earnings of people who moved with the earnings of people who did not move, it is necessary to identify not only those who moved to a new location, but also to differentiate those who remained in their new location from those who either moved again or returned to their original location. Thus, for this analysis, I restrict the definition of "movers" to those who relocated to a different county or state *then remained in that new location* for the ensuing 5 years.

Likewise, among those who did not move in a given year or years, it is necessary to differentiate those who remained in their original location from those who moved in later years. I thus define "nonmovers" as individuals who did not move in a given year of observation and continued to reside in the same county 5 years later.

Not all moves, particularly short-distance moves, result in a worker changing employers. Similarly, most changes of employer occur without requiring a move to a new county or state. Therefore, I examine trends in geographic mobility and employer change separately and in combination. Specifically, I explore the overall geographic mobility rate, then the overall employer-change rate. Then, among movers, I estimate the proportion who also changed employers; and among those who changed employers, I estimate the proportion who also moved.

The CWHS includes data on county and state of residence for all years since 1993 in which an individual had earnings.[5] I examine average rates of geographic mobility and employer change in three 3-year periods (1994–1996, 2004–2006, and 2014–2016), which span nearly the entire range of years for which data on county and state of residence are available in the CWHS. I compute annual earnings, rates of geographic mobility, and rates of employer change as 3-year annual averages because multiyear trends are less susceptible to statistical anomalies than are single-year data.

## *Geographic Mobility Rate and Employer Change Rate*

Charts 1 and 2 respectively show the annual average percentage of men and women who moved during each of the 3-year observation periods. The percentages are plotted for five age groups. In each age group, the plots show mobility rates in each 3-year period for men or women overall and for those in the lowest and highest quartiles of average annual earnings in the 3-year period ending with (and including) the year they moved. Because I observe events that occur over a 3-year period, some individuals do not remain in the same 5-year age group for the full period. For example, some men who were aged 25–29 in 1994 were in the 30–34 age group in 1996. However, these small changes in the composition of each age group during the observation periods had no material impact on either the descriptive statistics or the regression model results discussed later.

Chart 1 shows that, among all men (shown in red), the proportion who moved declined monotonically with age. In all three periods, men aged 25–29 were more than twice as likely as men aged 45–49 to have moved. From 1994–1996 to 2004–2006, there was little change in the rate of geographic mobility in any of the 5-year age groups; but from 2004–2006 to 2014–2016, geographic mobility fell among all age groups, possibly reflecting the lingering effects of the 2007–2009 recession (Partridge and others 2012; Goetz 2014).

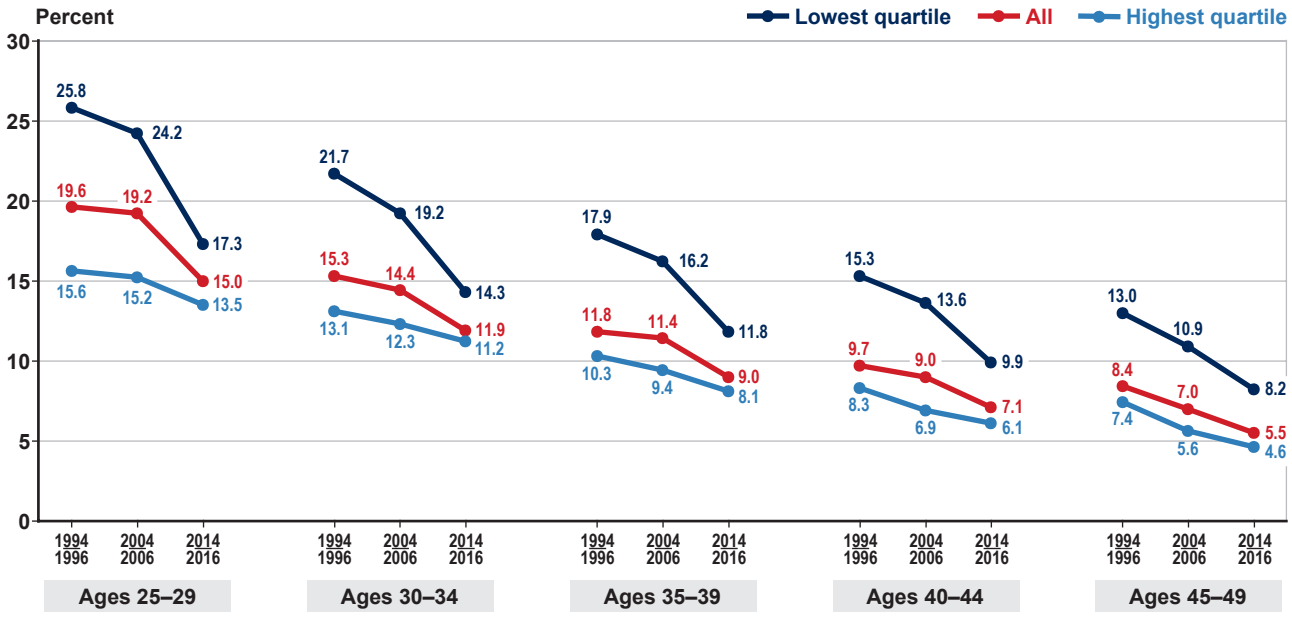
Geographic mobility rates among men in the lowest earnings quartile (shown in dark blue) declined between 1994–1996 and 2014–2016, but in all five age groups and in all three periods, annual geographic mobility rates were higher among men in the lowest earnings quartile than they were among all men in the same age group. Within each age group, the annual geographic mobility rates of men in the highest earnings quartile (shown in light blue) were lower than those of men overall, and were substantially lower than those of men in the lowest earnings quartile. As with the lowest earnings quartile, the percentage of men in the highest earnings quartile who moved declined between 1994–1996 and 2014–2016.

Chart 2 shows the annual average percentage of women who moved during each of the 3-year observation periods. Women were slightly less likely than men to have moved in each period. As was the case with men, the proportion of women who moved declined monotonically with age. Among all women, those aged 25–29 were almost three times more likely to have moved to a new county or state than those aged 45–49. From 1994–1996 to 2014–2016, the rate of geographic mobility declined in each of the 5-year age groups.

Among women whose 3-year average annual earnings placed them in the lowest earnings quartile for their age group, geographic mobility rates were higher than those of all women in the same age group. Similar to the trend among men, geographic mobility rates among women in the lowest earnings quartile declined between 1994–1996 and 2014–2016.

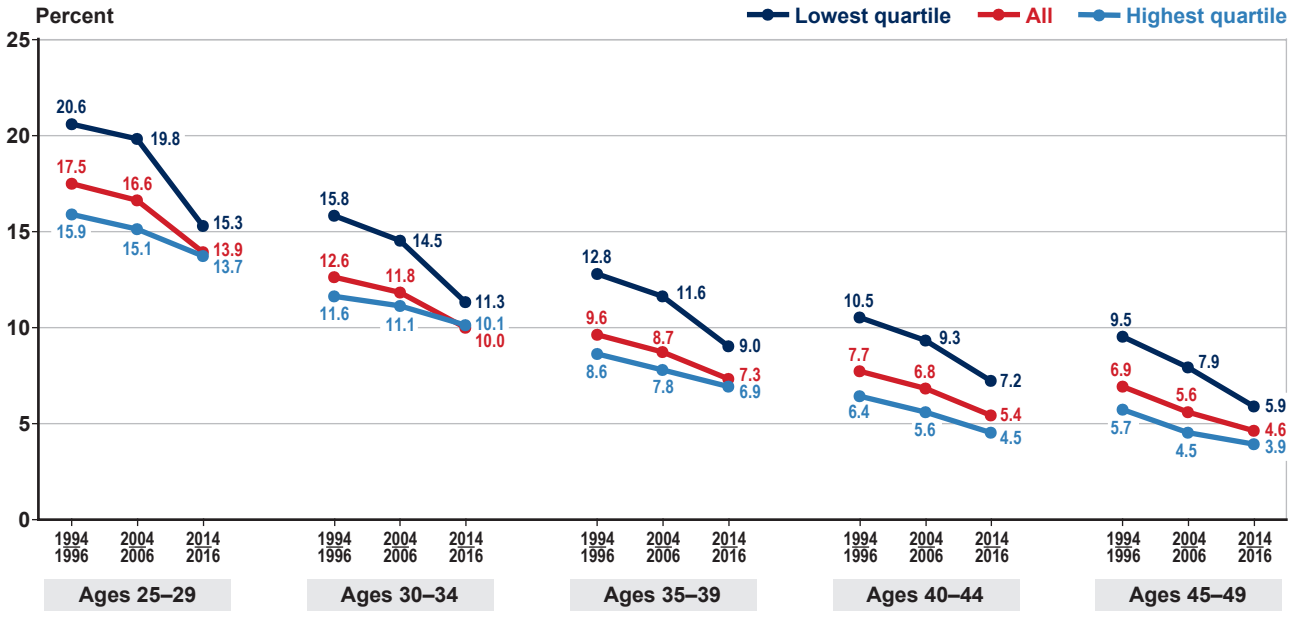
Within each age group, women in the highest earnings quartile had lower annual geographic mobility rates than did women overall and their mobility rates were substantially lower than those of women in the lowest earnings quartile. As was true of women in the lowest earnings quartile, the percentage of women in the highest quartile who moved declined between 1994–1996 and 2014–2016.

**Chart 1.**
**Men who moved to another county or state, by age group: Overall and by selected earnings quartile, various periods 1994–2016 (annual average percentages)**




SOURCE: Author's calculations using CWHS data.

**Chart 2.**
**Women who moved to another county or state, by age group: Overall and by selected earnings quartile, various periods 1994–2016 (annual average percentages)**




SOURCE: Author's calculations using CWHS data.

Chart 3 shows the annual average percentage of men who changed employers during the three observation periods. In this analysis, a change in the Employer Identification Number (EIN) recorded in the CWHS from one year to the next indicates a change of employer.[6] In all three periods, more men changed employers than moved to a new county or state. The average annual rate of employer change declined with age, but even among men aged 45–49, about one-sixth changed employers in a typical year. Unlike geographic mobility rates, the annual employer-change rates did not substantially decline. For example, among all men aged 25–29, the annual average percentage who changed employers decreased less than 1 percentage point from 1994–1996 to 2014–2016. Among all men aged 45–49, the annual average employer-change rate increased from 16.2 percent in 1994–1996 to 18.0 percent in 2014–2016.

Men in their age group's lowest earnings quartile were much more likely to have changed employers than all men within that age group. For example, in the period 2014–2016, among men aged 25–29 in the lowest earnings quartile, an annual average of 54.8 percent changed employers, compared with 36.4 percent of all men in that age group. Among men aged 45–49 in the lowest earnings quartile, an annual average of 32.6 percent changed employers in the period 2014–2016, compared with 18.0 percent of men aged 45–49 overall.

In each age group, men in the highest earnings quartile were less likely to have changed employers than men overall. For example, 21.5 percent of men aged 25–29 in the highest earnings quartile changed employers annually in the period 2014–2016, compared with 36.4 percent of all men aged 25–29. Among men aged 45–49, 11.9 percent of those in the highest earnings quartile changed employers in the period 2014–2016, compared with 18.0 percent of all men aged 45–49.

Average annual rates of employer change did not decline among men in the highest earnings quartile between 1994–1996 and 2014–2016. Among men in the 25–29 and 30–34 age groups, the proportion who changed employers rose slightly over time. Among the three older age groups, the annual average proportion of men in the highest earnings quartile who changed employers was essentially the same in 2014–2016 as it had been in 1994–1996.

Chart 4 shows the annual average percentage of women who changed employers during the three observation periods. As with men, women changed employers more often than they moved to a new county or state. In each 3-year period, an annual average of about 36 percent of all women aged 25–29 changed employers. The annual average rate of employer change declined with age, but even among all women aged 45–49, about one-sixth changed employers in a typical year.

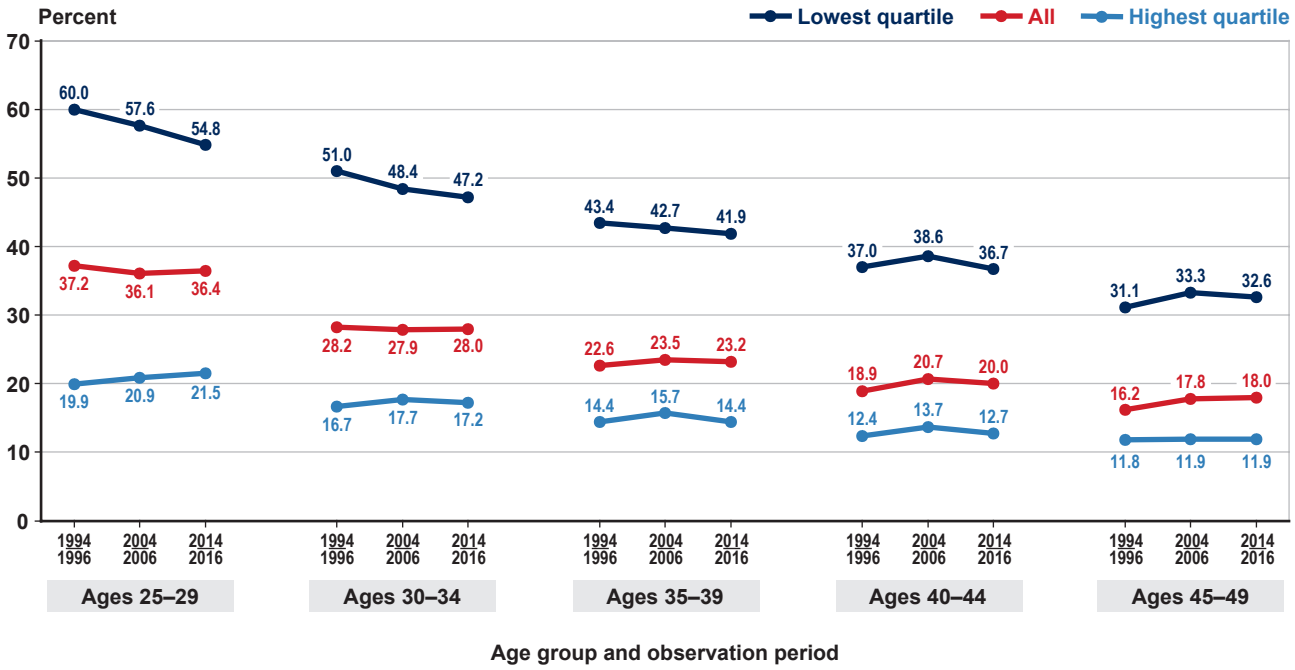
Unlike geographic mobility rates, average annual employer-change rates did not decline among women between 1994–1996 and 2014–2016. Of all women aged 25–29, an annual average of 36.4 percent changed employers in 2014–2016, up slightly from 35.6 percent in 1994–1996. Among all women aged 45–49, the annual average rate of employer change rose from 16.4 percent in 1994–1996 to 17.4 percent in 2014–2016.

In all three periods, women in the lowest earnings quartile for their age group were much more likely to have changed employers than were all women in that age group. Among women aged 25–29 in the lowest earnings quartile, an annual average of 51.3 percent changed employers during 2014–2016, compared with 36.4 percent of all women aged 25–29. Among women aged 45–49 in the lowest earnings quartile, an average of 28.4 percent changed employers each year during 2014–2016, compared with 17.4 percent of all women aged 45–49.

From 1994–1996 to 2014–2016, annual average employer-change rates declined for women in the lowest earnings quartile of each of the age groups younger than 40. Among women in the 40–44 and 45–49 age groups, the annual average proportion who changed employers was approximately the same in 2014–2016 as it had been in 1994–1996.

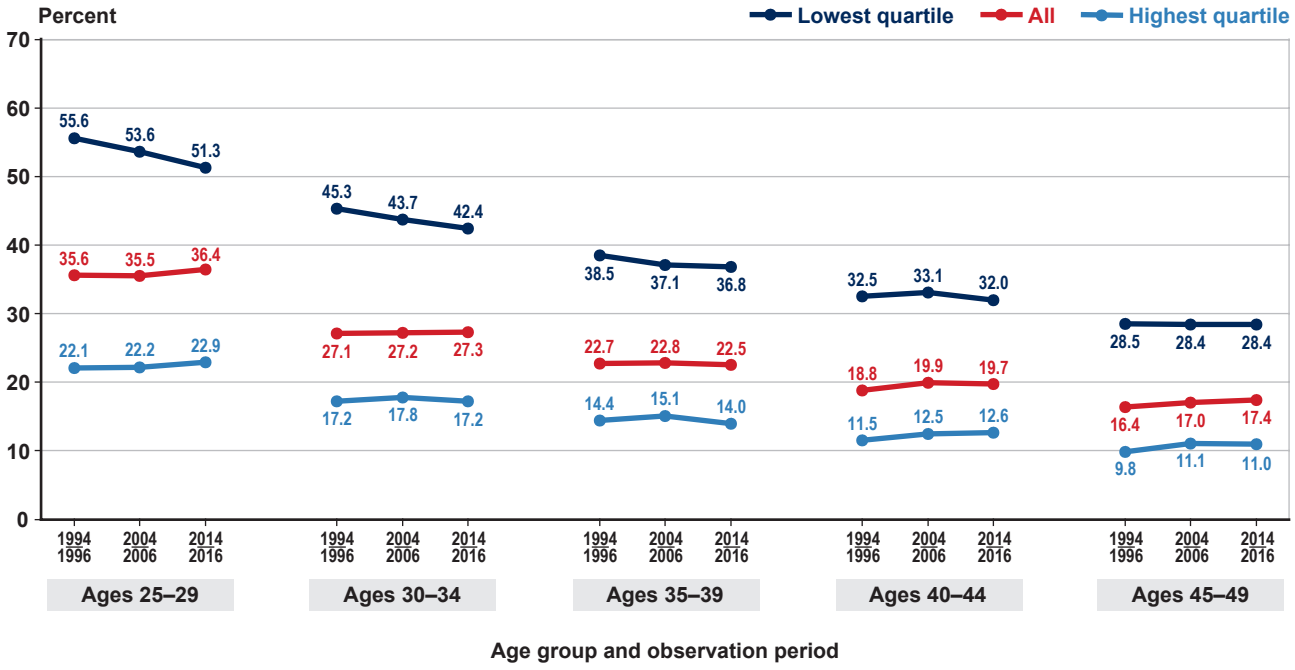
Women in the highest earnings quartile of their age groups were much less likely to have changed employers than were all women of the same age. In the highest earnings quartile of the 25–29 age group, 22.9 percent changed employers annually in the period 2014–2016, compared with 36.4 percent of all women aged 25–29. In the highest earnings quartile of the 45–49 age group, 11.0 percent of women changed employers annually in the period 2014–2016, compared with 17.4 percent of all women aged 45–49. With the slight exception of the 35–39 age group, average annual employer-change rates did not decline among women in the highest earnings

**Chart 3.**
**Men who changed employers, by age group: Overall and by selected earnings quartile, various periods 1994–2016 (annual average percentages)**




SOURCE: Author's calculations using CWHS data.

**Chart 4.**
**Women who changed employers, by age group: Overall and by selected earnings quartile, various periods 1994–2016 (annual average percentages)**




SOURCE: Author's calculations using CWHS data.

quartile between 1994–1996 and 2014–2016. Among women aged 25–29, 22.9 percent changed employers in 2014–2016, compared with 22.1 percent in 1994–1996. Among those aged 45–49, 11.0 percent changed employers in 2014–2016, compared with 9.8 percent in 1994–1996.

Moving to a new county or state often involves changing employers, but changing employers does not as often require moving to a new location. Chart 5 shows, for men who moved during one of the three observation periods, the percentages who also changed employers in that period (shown in blue). In each period, almost two-thirds of men aged 25–29 who moved also changed employers, and about half of men aged 45–49 who moved also changed employers. Among men in all age groups except 45–49, the proportion of movers who also changed employers declined slightly from 1994–1996 through 2014–2016. The proportion of movers who also changed employers would be higher except that some moves across county or state lines occur within a single commuting zone and therefore are less likely to involve a change of employers. For example, the Washington, DC commuting zone includes the District of Columbia, five counties in Maryland, and six counties in Virginia, making relocations across county or state lines without changing employers feasible for many workers there.

Chart 5 also shows, for men who changed employers, the percentages who also moved in the same observation period (shown in red). Although a majority of workers who move also change employers, most people who change employers do so without moving; and among men who changed employers, the proportion who also moved declined sharply from 1994–1996 to 2014–2016. Among men aged 25–29 who changed employers, the proportion who also moved declined from 35.2 percent in 1994–1996 to 26.0 percent in 2014–2016. Among men aged 45–49 who changed employers, the proportion who also moved declined from 25.9 percent in 1994–1996 to 15.6 percent in 2014–2016.

Chart 6 repeats Chart 5 for women. In each period and in all age categories, among women who moved, the proportion who also changed employers was similar to the proportion among the corresponding age group of men. In all three periods, among women aged 25–29 who moved, more than 60 percent also changed employers. Among women aged 45–49 who moved, about half also changed employers.

Among women who changed employers, the proportions who also moved were similar to the proportions among the corresponding age groups of men. From 1994–1996 to 2014–2016, the annual average percentage of employer-changing women who also moved declined sharply. Among women aged 25–29, this proportion declined from 31.3 percent in 1994–1996 to 24.5 percent in 2014–2016. For those aged 45–49, the proportion in 2014–2016 was 13.1 percent, down from 21.0 percent in 1994–1996.

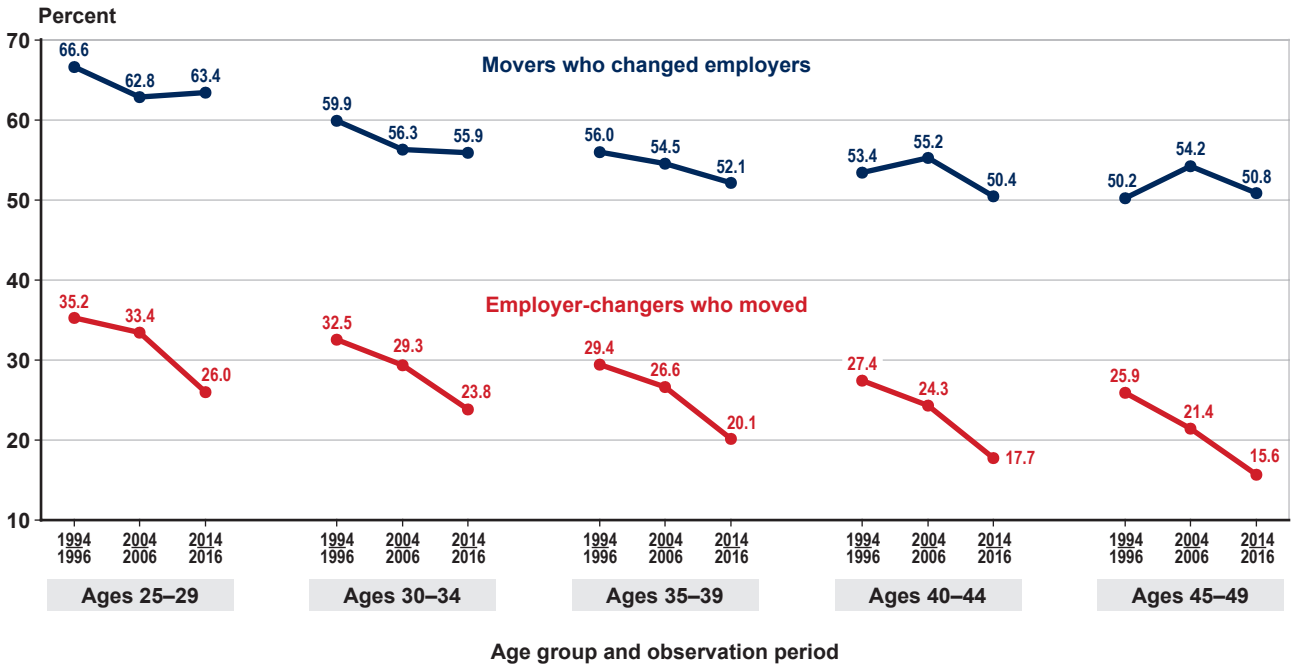
In summary, Charts 1 through 6 show that the average annual proportion of men and women aged 25–49 who moved to a new county or state declined substantially from 1994–1996 through 2014–2016. The proportion of men and women who changed employers, however, changed relatively little. The latter finding contrasts with some earlier studies that detected a downward trend in rates of employer change among American workers. Among men who moved, the proportion who also changed employers declined slightly over the period from 1994–1996 through 2014–2016. Among men who changed employers over that period, the proportion who also moved fell sharply. For women who moved, there was relatively little change in the proportion who also changed employers over the period from 1994–1996 through 2014–2016. For women who changed employers, the proportion who also moved declined substantially, mirroring the trend among men.

## *Multivariate Analysis of Geographic Mobility and Employer Change*

This section discusses the results of several regression models that test the statistical relationship between a range of individual and geographic variables and the probability that an individual moved or changed employers. Table 1 shows the results for two logistic regressions: In model 1, the dependent variable indicates whether the individual moved to a new county or state in the previous calendar year, and in model 2, the dependent variable indicates whether the individual changed employers in the previous calendar year. Both models control for the 3-year observation period, age, race, foreign or domestic birthplace, region of residence, and whether the county of residence was metropolitan or nonmetropolitan.[7] Region and county are defined as the place of residence in the year before an individual moved or changed employer, or in the same year for a member of the comparison group who did not.

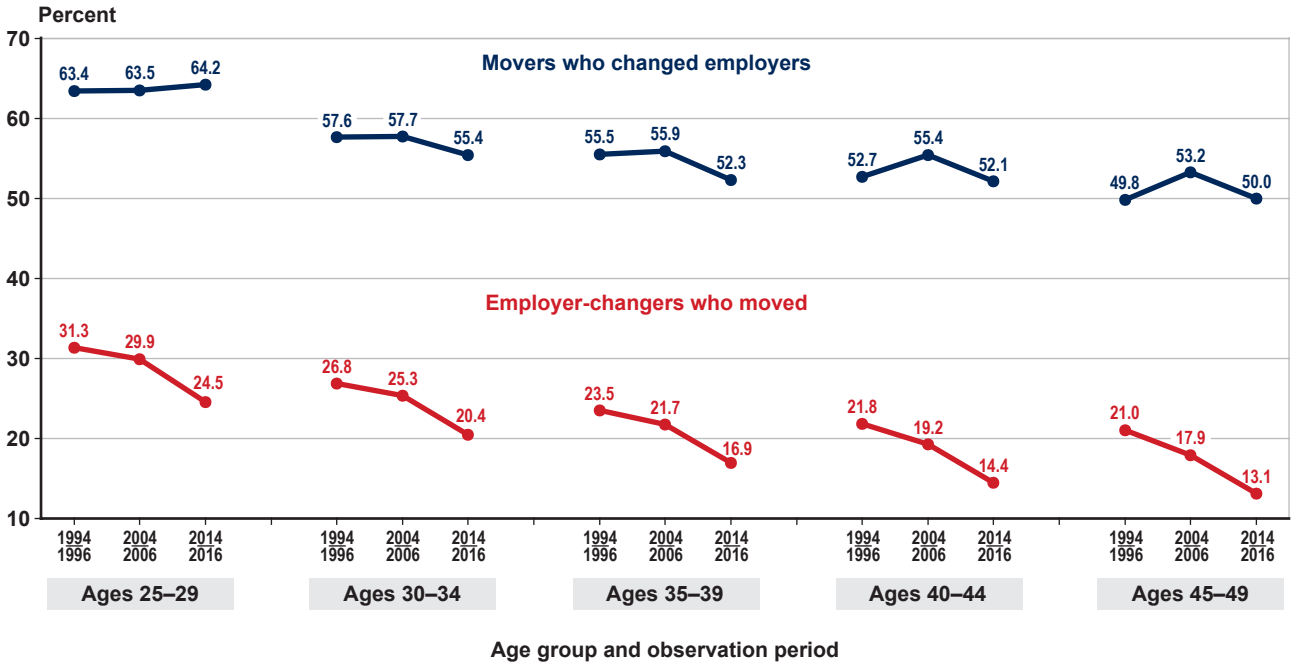
The main economic variables of interest in each regression are the quartile rank of each person's mean annual earnings in the 3 years before the year in which

**Chart 5.**
**Interactions between moving and changing employers among men, by age group: Percentage of movers who changed employers, and percentage of employer-changers who moved, various periods 1994–2016**




SOURCE: Author's calculations using CWHS data.

**Chart 6.**
**Interactions between moving and changing employers among women, by age group: Percentage of movers who changed employers, and percentage of employer-changers who moved, various periods 1994–2016**




SOURCE: Author's calculations using CWHS data.

**Table 1.**
**Relationship of selected characteristics to the probability of having moved or changed employers in the past year among workers aged 25–49, by sex: Logistic regression results**

| Independent variable | Model 1: Moved in past year | | Model 2: Changed employer in past year | |
|---|---|---|---|---|
| | Marginal effect [a] | Standard error | Marginal effect [a] | Standard error |
| | | ***Men*** [b] | | |
| Observation period | | | | |
| 2004–2006 | -0.0047* | 0.0022 | 0.0051* | 0.0017 |
| 2014–2016 | -0.0300* | 0.0141 | 0.0077* | 0.0026 |
| Foreign place of birth | 0.0057* | 0.0027 | -0.0234* | 0.0080 |
| Age (1-year increment) | -0.0037* | 0.0017 | -0.0038* | 0.0013 |
| White, non-Hispanic | 0.0092* | 0.0043 | -0.0114* | 0.0039 |
| Region | | | | |
| Midwest | 0.0046* | 0.0022 | 0.0025* | 0.0009 |
| South | 0.0218* | 0.0103 | 0.0249* | 0.0085 |
| West | 0.0034* | 0.0016 | 0.0201* | 0.0069 |
| Metropolitan county | -0.0697* | 0.0327 | -0.0035* | 0.0012 |
| Quartile of mean annual earnings [c] | | | | |
| 4th (highest) | -0.0599* | 0.0282 | -0.2541* | 0.0867 |
| 3rd | -0.0579* | 0.0272 | -0.2253* | 0.0769 |
| 2nd | -0.0379* | 0.0178 | -0.1411* | 0.0482 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | -0.0154* | 0.0073 | -0.0030* | 0.0010 |
| Median household income | 0.0499* | 0.0235 | 0.0396* | 0.0135 |
| | | ***Women*** [d] | | |
| Observation period | | | | |
| 2004–2006 | -0.0050* | 0.0025 | 0.0022* | 0.0007 |
| 2014–2016 | -0.0236* | 0.0117 | 0.0048* | 0.0016 |
| Foreign place of birth | 0.0013* | 0.0006 | -0.0249* | 0.0084 |
| Age (1-year increment) | -0.0038* | 0.0019 | -0.0055* | 0.0018 |
| White, non-Hispanic | 0.0072* | 0.0036 | -0.0202* | 0.0068 |
| Region | | | | |
| Midwest | 0.0020* | 0.0010 | 0.0093* | 0.0031 |
| South | 0.0189* | 0.0094 | 0.0243* | 0.0082 |
| West | 0.0049* | 0.0025 | 0.0219* | 0.0074 |
| Metropolitan county | -0.0640* | 0.0319 | 0.0018* | 0.0006 |
| Quartile of mean annual earnings [c] | | | | |
| 4th (highest) | -0.0340* | 0.0169 | -0.2135* | 0.0719 |
| 3rd | -0.0321* | 0.0160 | -0.1756* | 0.0591 |
| 2nd | -0.0208* | 0.0104 | -0.0967* | 0.0326 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | -0.0140* | 0.0070 | -0.0166* | 0.0056 |
| Median household income | 0.0447* | 0.0223 | 0.0228* | 0.0077 |

SOURCE: Author's calculations using CWHS data.

NOTE: * = statistically significant at the 0.01 level.

a. The change in the probability of the event represented by the dependent variable, either relative to the omitted categorical independent variable or in response to a one-unit change in a continuous independent variable, averaged across all observations in the sample.

b. Model 1: 3,759,864 observations; log likelihood = -1,208,341; Hosmer-Lemeshow test $\chi^2$ = 248.4; probability > $\chi^2$: <.0001.
Model 2: 3,762,224 observations; log likelihood = -1,857,042; Hosmer-Lemeshow test $\chi^2$ = 10,009.5; probability > $\chi^2$: <.0001.

c. In the 3-year observation period.

d. Model 1: 3,656,583 observations; log likelihood = -1,044,569; Hosmer-Lemeshow test $\chi^2$ = 855.6; probability > $\chi^2$: <.0001.
Model 2: 3,657,489 observations; log likelihood = -1,805,963; Hosmer-Lemeshow test $\chi^2$ = 2,628.0; probability > $\chi^2$: <.0001.

a move or employer change did or did not occur; and two indicators of local economic conditions. These economic indicators are the ratio of the 3-year average unemployment rate in the person's county of residence to the national 3-year average unemployment rate and the ratio of median household income in the person's county of residence to national median household income. For both men and women, 3-year mean annual earnings were ranked by quartile in each of the five age groups from 25–29 through 45–49. The county/national unemployment-rate ratio was based on 3-year averages computed from the Current Population Survey. For the period 1994–1996, the county/national household income ratio was based on data from the 1990 census. For the two later periods, the income ratios were based on data from the American Community Survey.

Table 1 shows the results of the two logistic regression models in which the samples consist of men and women aged 25–49. For each independent variable, the table shows the average marginal effect (with an indicator of statistical significance) and the standard error. The sample for each regression represents approximately 3.7 million observations over three 3-year periods.[8] In this sample, annual averages of 10.7 percent of men moved and 23.0 percent changed employers; the corresponding percentages for women are 8.9 percent and 22.0 percent.

In model 1, the average marginal effect represents the change in the probability of having moved to a new county or state either in response to a one-unit change in an independent variable or relative to the omitted reference variable, averaged across all observations in the sample. The marginal effects of the variables representing 2004–2006 and 2014–2016 were negative and statistically significant relative to 1994–1996, other things being equal. Among the other independent variables, men born outside the United States were slightly more likely to have moved than native-born men were. Non-Hispanic white men were more likely to have moved than other men. The probability of moving declined with age. Men residing in the Midwest, South, or West were more likely to have moved than men residing in the Northeast, and men who lived in metropolitan-area counties were less likely to have moved than men who lived in nonmetropolitan counties.

Chart 1 showed that men in the lowest earnings quartile for their age group had higher annual geographic mobility rates than did men in the top quartile. The same relationship is present in the regression results. Relative to men in the first (lowest) earnings quartile for their 5-year age group, the annual probability of moving was 3.8 percentage points lower for men in the second earnings quartile. For men in the third and fourth (highest) earnings quartiles, the annual probabilities of moving were 5.8 percentage points and 6.0 percentage points lower, respectively, than for men in the first earnings quartile.

The variables representing local economic conditions also had statistically significant relationships with the likelihood of moving. The probability of moving was 1.5 percentage points lower for men who resided in counties with local-to-national unemployment-rate ratios greater than 1 than that for men in counties with lower ratios. The probability of moving was 5.0 percentage points higher for men who resided in counties with higher than average local-to-national median household income ratios, all else being equal. Thus, although the probability of moving was negatively correlated with successively higher individual earnings quartiles, it was positively correlated with county median household income. Regressions run separately on men in each earnings quartile showed an average marginal effect for county median household income of 0.101 in the lowest earnings quartile compared with an average marginal effect of just 0.016 in the highest earnings quartile (not shown). This suggests that the characteristics of high-income counties, such as higher average educational attainment, may promote greater geographic mobility for lower earners in those counties, but that the effect dissipates as one's own earnings rise.[9]

In model 2, the average marginal effect represents the change in the probability of having changed employers either in response to a one-unit change in an independent variable or relative to the omitted reference variable, averaged across all observations in the sample. In this model, the average marginal effects of the variables representing the years 2004–2006 and 2014–2016 were small but positive for men and women alike. As Chart 3 showed, unlike geographic mobility, rates of employer change among men did not decline over time. In the regression analysis, men born outside the United States were less likely to have changed employers than were native-born men. Men who reported their race/ethnicity as non-Hispanic white were less likely to have changed employers than other men were. The probability of changing employers declined with age. Men in the Midwest, South, and West were more likely to have changed employers than those in the Northeast, and those in metropolitan

counties were slightly less likely to have changed employers than were men in nonmetropolitan counties.

Relative to men in the lowest earnings quartile, those in the other three quartiles were substantially less likely to have changed employers, all else being equal. Men in the fourth (highest) earnings quartile were 25 percentage points less likely to change employers than those in the first (lowest) quartile. The probabilities of employer change in the third and second quartiles were 23 percentage points and 14 percentage points lower, respectively, than those of men in the lowest earnings quartile. Residing in a county with a local-to-national average annual unemployment-rate ratio of 1 or more had a small but statistically significant negative correlation with the annual probability of changing employers.

For men who resided in counties in which median household income exceeded the national median household income, the probability of changing employers was 4.0 percentage points higher than average, all else being equal. Thus, although the probability of changing employers was negatively correlated with successively higher earnings quartiles, it rose with county median household income—the same pattern as that for the annual probability of moving. Further mirroring the results for geographic mobility, regressions run separately on men in each earnings quartile showed that the average marginal effect of county median household income was substantially larger for men in the lowest earnings quartile than for men in higher quartiles (not shown). This suggests that employer change may be easier for low earners in high-income counties, with the effect dissipating as one's own earnings rise.

For women, the average annual probability of moving had the same signs as those of men for the individual earnings quartile, county unemployment rate, and county median household income variables, but the average marginal effects for women were slightly smaller. For the average annual probability of changing employers, the signs for the three economic variables also were the same for both men and women. The average marginal effect of the county unemployment rate was slightly larger for women and those of the other economic variables were slightly smaller for women.

As Charts 5 and 6 illustrated, approximately two-thirds of men and women who moved to another county or state also changed employers, and about one-third of those who changed employers also moved to another county or state. By constructing two subsamples—one comprising individuals who changed employers and the other consisting of those who did not—and running the same logistic models described above separately on each subsample, we can examine the statistical relationship of selected personal and geographic traits to geographic mobility, conditional on having changed or not changed employers. Table 2 shows the results of logistic regressions run separately on men and women who changed employers in the previous year, and men and women who did not change employers.

Table 2, subsample 1 shows that, among men who changed employers in the preceding year, the independent variables representing earnings quartile have the same sign and approximately the same magnitude as in the regression run on the full sample of men (Table 1, model 1). The average marginal effects for high county unemployment rate and median household income, however, are larger for the subsample who changed employers. Other factors being equal, residing in a high-unemployment county had a stronger negative correlation with geographic mobility among men who changed employers than among the full sample. This may support the hypothesis that workers in economically disadvantaged areas are relatively less able to migrate to areas with better employment opportunities (Raphael and Riker 1999; Foster 2017). Similarly, the positive marginal effect associated with high county median household income was larger for the subsample of men who changed employers than for the sample as a whole. This could indicate that characteristics of higher-income counties, such as higher average educational attainment, promote greater geographic mobility.

For women who changed employers in the preceding year, each independent variable in the Table 2 regression estimating the probability of moving to a new county or state had the same sign as that for men, but the estimated average marginal effects were smaller in most cases.

## *Annual Earnings of Movers and Nonmovers*

Data from the CWHS in Charts 1–4 show that annual geographic mobility rates declined substantially among both men and women aged 25–49 from 1994–1996 through 2014–2016. Annual rates of employer change remained relatively stable over that period among both men and women, except for those younger than 40 in the lowest earnings quartile, for whom employer change declined. Kennan and Walker (2011) found that the prospect of higher earnings in other locations is a significant incentive for geographic

**Table 2.**
**Relationship of selected characteristics to the probability of having moved in the past year among workers aged 25–49, by sex and employer-change status: Logistic regression results**

| Independent variable | Subsample 1: Changed employer in past year | | Subsample 2: Did not change employer in past year | |
|---|---|---|---|---|
| | Marginal effect [a] | Standard error | Marginal effect [a] | Standard error |
| | | ***Men*** [b] | | |
| Observation period | | | | |
| 2004–2006 | -0.0184* | 0.0040 | -0.0015* | 0.0007 |
| 2014–2016 | -0.0830* | 0.0182 | -0.0149* | 0.0067 |
| Foreign place of birth | 0.0216* | 0.0047 | 0.0067* | 0.0030 |
| Age (1-year increment) | -0.0045* | 0.0010 | -0.0026* | 0.0012 |
| White, non-Hispanic | 0.0292* | 0.0064 | 0.0059* | 0.0027 |
| Region | | | | |
| Midwest | 0.0059* | 0.0013 | 0.0036* | 0.0016 |
| South | 0.0316* | 0.0069 | 0.0132* | 0.0060 |
| West | 0.0026 | 0.0006 | -0.0012* | 0.0005 |
| Metropolitan county | -0.1534* | 0.0337 | -0.0437* | 0.0198 |
| Quartile of mean annual earnings [c] | | | | |
| 4th (highest) | -0.0616* | 0.0135 | 0.0038* | 0.0017 |
| 3rd | -0.0499* | 0.0110 | -0.0002 | 0.0001 |
| 2nd | -0.0356* | 0.0078 | 0.0028* | 0.0013 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | -0.0430* | 0.0094 | -0.0065* | 0.0029 |
| Median household income | 0.0705* | 0.0155 | 0.0330* | 0.0149 |
| | | ***Women*** [d] | | |
| Observation period | | | | |
| 2004–2006 | -0.0092* | 0.0023 | -0.0038* | 0.0019 |
| 2014–2016 | -0.0589* | 0.0149 | -0.0133* | 0.0066 |
| Foreign place of birth | 0.0134* | 0.0034 | 0.0028* | 0.0014 |
| Age (1-year increment) | -0.0052* | 0.0013 | -0.0022* | 0.0011 |
| White, non-Hispanic | 0.0335* | 0.0085 | 0.0034* | 0.0017 |
| Region | | | | |
| Midwest | 0.0006 | 0.0001 | 0.0010 | 0.0005 |
| South | 0.0329* | 0.0083 | 0.0101* | 0.0050 |
| West | 0.0116* | 0.0029 | -0.0015* | 0.0007 |
| Metropolitan county | -0.1456* | 0.0368 | -0.0407* | 0.0201 |
| Quartile of mean annual earnings [c] | | | | |
| 4th (highest) | -0.0259* | 0.0065 | 0.0089* | 0.0044 |
| 3rd | -0.0203* | 0.0051 | 0.0057* | 0.0028 |
| 2nd | -0.0162* | 0.0041 | 0.0036* | 0.0018 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | -0.0328* | 0.0083 | -0.0053* | 0.0026 |
| Median household income | 0.0639* | 0.0161 | 0.0327* | 0.0161 |

SOURCE: Author's calculations using CWHS data.

NOTE: * = statistically significant at the 0.01 level.

a. The change in the probability of the event represented by the dependent variable, either relative to the omitted categorical independent variable or in response to a one-unit change in a continuous independent variable, averaged across all observations in the sample.

b. Subsample 1: 2,897,620 observations; log likelihood = -655,974; Hosmer-Lemeshow test $\chi^2$ = 287.2; probability > $\chi^2$: <.0001.
Subsample 2: 862,244 observations; log likelihood = -469,430; Hosmer-Lemeshow test $\chi^2$ = 584.3; probability > $\chi^2$: <.0001.

c. In the 3-year observation period.

d. Subsample 1: 2,842,852 observations; log likelihood = -558,955; Hosmer-Lemeshow test $\chi^2$ = 963.5; probability > $\chi^2$: <.0001.
Subsample 2: 813,731 observations; log likelihood = -412,292; Hosmer-Lemeshow test $\chi^2$ = 358.3; probability > $\chi^2$: <.0001.

mobility. The charts in this section compare the annual earnings of men and women who moved to a new county or state with the earnings of those who did not move. Specifically, the charts show, for workers who moved in 1994–1996 or 2004–2006, average annual earnings in the 3-year period before moving and in the 4th through 6th years after moving, and compare them with the earnings in the same years of people who did not move. In this section, "earnings" refers to the median value of the 3-year annual average earnings for the members of a given age group in an observation period.

Chart 7 shows earnings among men in each of the five age groups for the periods 1994–1996 and 2004–2006. For men who moved, the chart shows earnings in the 3-year period up to and including the year of the move. Some methodological points bear repeating here: To provide meaningful comparisons, the earnings of men who did not move are shown for the same years. The sample includes only men who remained for 5 years in one residence—in either the mover's new location or the nonmover's same location. The sample thus excludes movers who returned, or who moved more than once in 5 years.

The earnings of men who moved in 1994–1996 were lower than those of nonmovers in all five age groups. Earnings among men aged 25–29 who moved during 1994–1996 were $24,603, or $3,233 (11.6 percent) less than those of nonmovers in the same age group ($27,836). Among men aged 45–49, earnings among movers were $59,786, or $5,046 (7.8 percent) less than similarly aged nonmovers ($64,832). Ten years later, the pattern persisted. Men aged 25–29 who moved during 2004–2006 had earnings ($28,079) that were $2,265 (7.5 percent) lower than those of nonmovers ($30,344). Among men aged 45–49, earnings among movers were $58,764, or $4,975 (7.8 percent) less than those of nonmovers ($63,739).

Chart 8 shows that for women, the relative earnings of movers and nonmovers were similar to those for men in 1994–1996, but differed slightly in 2004–2006. Earnings of women aged 25–29 who moved during 1994–1996 were $21,947 in the 3 years up to and including the year they moved, or $774 (3.4 percent) less than the earnings of nonmovers in the same period ($22,721). Earnings among women aged 45–49 were $35,045 for movers, or $3,116 (8.2 percent) less than those of nonmovers ($38,161). Ten years later, however, the pattern differed. During 2004–2006, the earnings of movers were lower than those of nonmovers for women in three of the five age groups, but were higher in the other two. For women in the 25–29, 40–44, and 45–49 age groups, the earnings of movers were lower than those of nonmovers. For women in the 30–34 and 35–39 age groups, the earnings of movers were higher than those of nonmovers.

Deciding whether to move to a new location might be influenced both by recent past earnings and by expectations about future earnings. Relatively low recent past earnings may prompt some workers to consider the possibility of earning higher wages elsewhere. In both 1994–1996 and 2004–2006, men's recent earnings were lower among those who moved than they were among men who, in the same period, did not. Among women, this relationship was also present in all five age groups in 1994–1996 and in three of the age groups in 2004–2006.

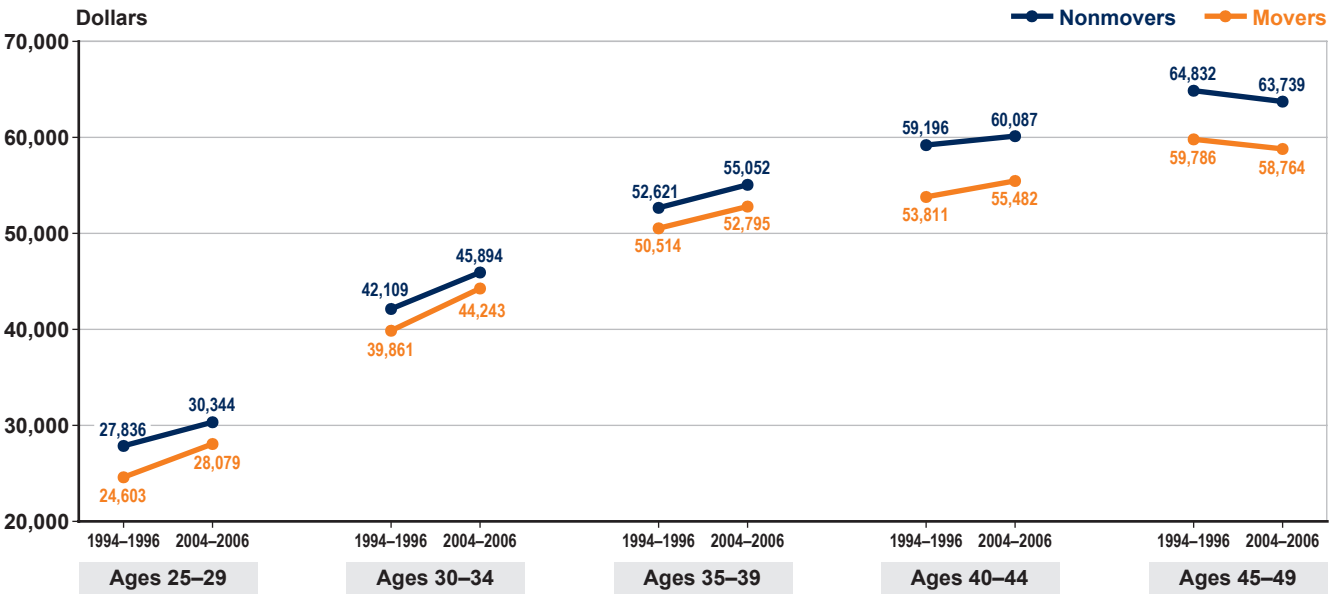
Chart 9 shows earnings in the 4th through 6th years after relocating among men who moved and in the same years among men who did not move. In both periods, men younger than 40 who moved had higher earnings 4–6 years after moving than their counterparts who did not move, even though movers had had lower earnings before they moved. Also in both periods, men aged 40–49 who moved had *lower* earnings 4–6 years after moving than men who did not move.

Among men aged 25–29 who moved during 1994–1996, postmove earnings were $51,606, or $4,691 (10.0 percent) higher than the earnings among nonmovers ($46,915). For men aged 45–49 who moved during 1994–1996, postmove earnings were $62,901, or $3,664 (5.5 percent) *lower* than the earnings of nonmovers ($66,565).

Ten years later, earnings had fallen for both movers and nonmovers, reflecting in part the effect of the Great Recession of 2007–2009. Among movers aged 25–29 during 2004–2006, postmove earnings were $48,813, or $5,396 (12.4 percent) higher than those of nonmovers ($43,417). Among men aged 45–49 who moved during 2004–2006, postmove earnings were $59,061, or $3,488 (5.6 percent) *lower* than those of nonmovers ($62,549).

Chart 10 repeats Chart 9 for women. In both periods, women aged 25–34 who moved had higher postmove earnings than women the same age in the same period who did not move, even though movers had had lower earnings before they moved. Also in both periods, women aged 40–49 who moved had *lower* postmove earnings than women the same age in the same period who did not move. Among women aged 35–39, the earnings of movers and nonmovers

## Chart 7.
**Men's earnings in the 3-year period ending with the year of relocation for movers, and in the same period for nonmovers, by age group, 1994–1996 and 2004–2006**

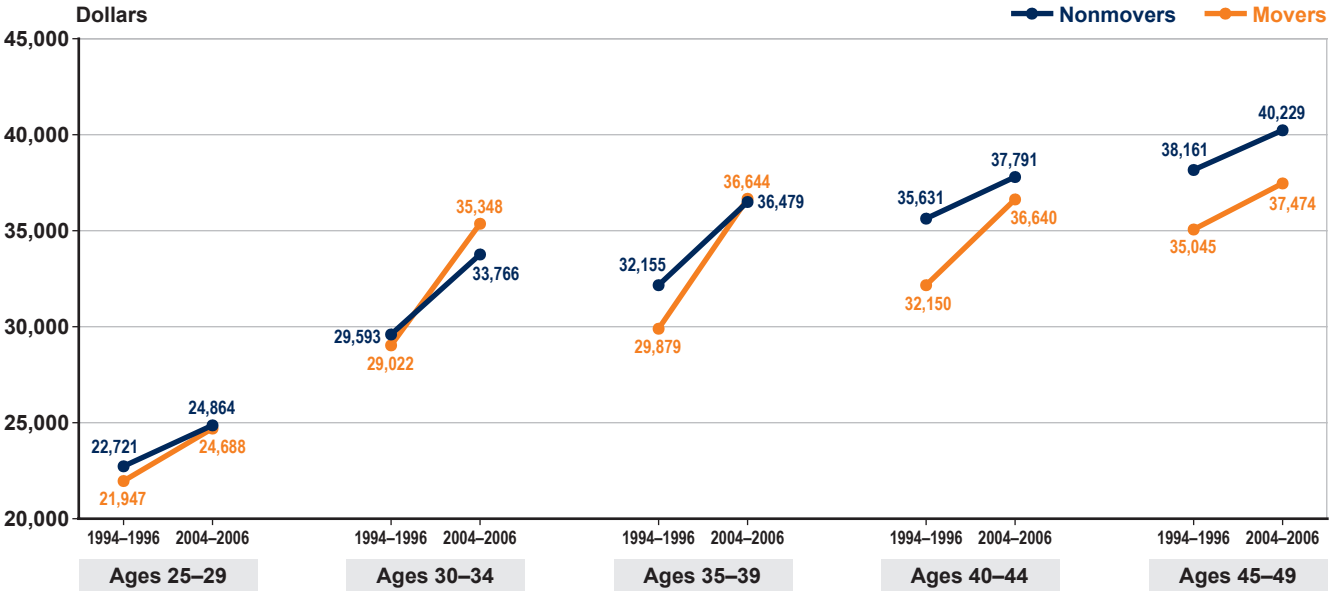



SOURCE: Author's calculations using CWHS data.

NOTES: Earnings are the medians of 3-year annual averages, expressed in 2016 dollars.

Earnings occurred in or prior to the move/nonmove observation period.

## Chart 8.
**Women's earnings in the 3-year period ending with the year of relocation for movers, and in the same period for nonmovers, by age group, 1994–1996 and 2004–2006**

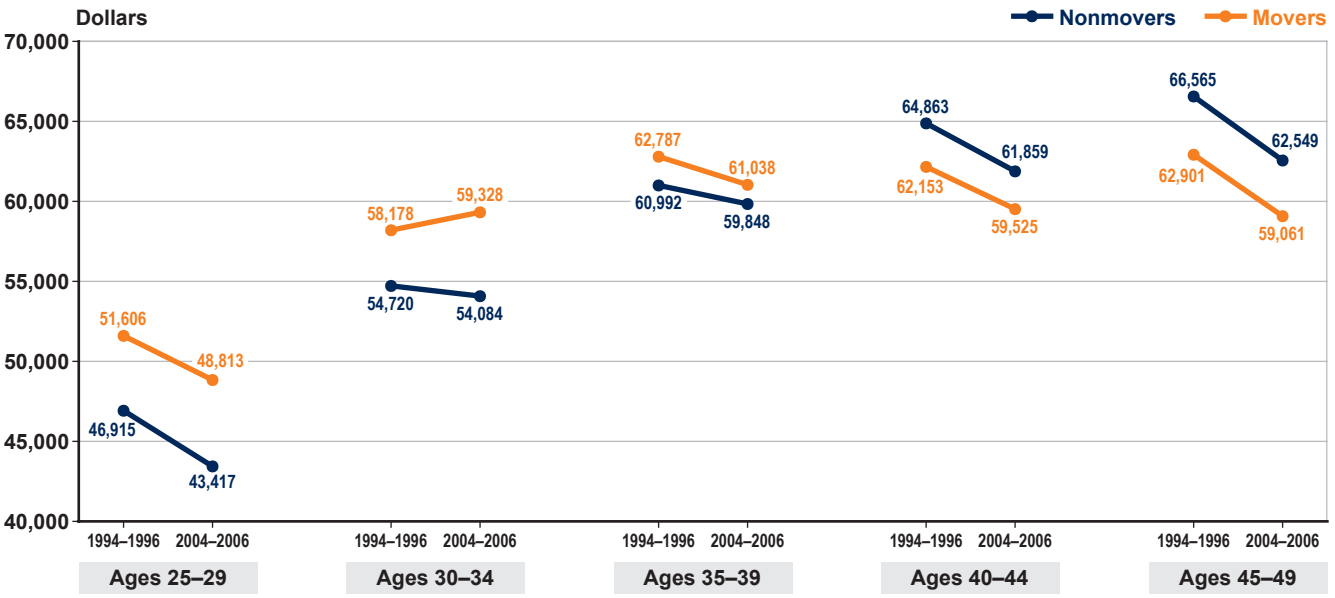



SOURCE: Author's calculations using CWHS data.

NOTE: Earnings are the medians of 3-year annual averages, expressed in 2016 dollars.

Earnings occurred in or prior to the move/nonmove observation period.

## Chart 9.
## Men's earnings in the 4th through 6th years after relocating for movers, and in the same period for nonmovers, by age group, 1994–1996 and 2004–2006

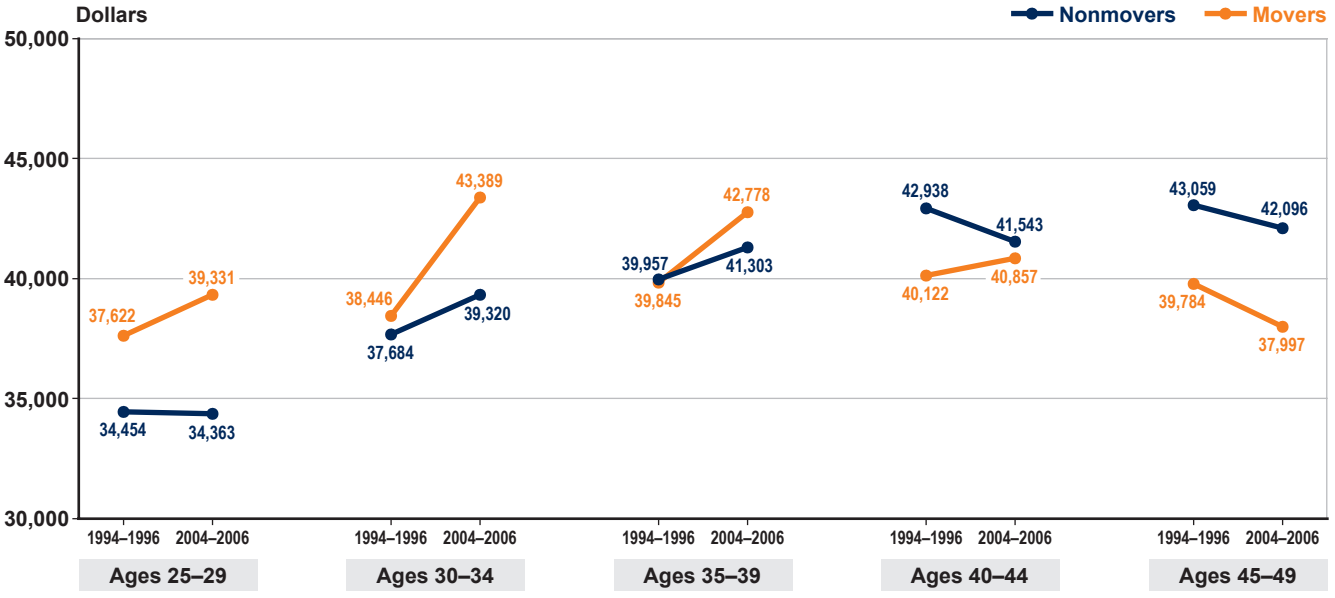



SOURCE: Author's calculations using CWHS data.

NOTES: Earnings are the medians of 3-year annual averages, expressed in 2016 dollars.

Earnings occurred in years following the move/nonmove observation period.

## Chart 10.
## Women's earnings in the 4th through 6th years after relocating for movers, and in the same period for nonmovers, by age group, 1994–1996 and 2004–2006




SOURCE: Author's calculations using CWHS data.

NOTE: Earnings are the medians of 3-year annual averages, expressed in 2016 dollars.

Earnings occurred in years following the move/nonmove observation period.

were about equal 4–6 years after the 1994–1996 period, and the earnings of movers were higher 4–6 years after the 2004–2006 period.

For women aged 25–29 who moved during 1994–1996, postmove earnings were $37,622, or $3,168 (9.2 percent) higher than those of nonmovers ($34,454). Among women aged 45–49 who moved in those years, postmove earnings were $39,784, or $3,275 (7.6 percent) *lower* than the earnings of nonmovers ($43,059).

Ten years later, the pattern was similar. Among women aged 25–29, the postmove earnings of those who moved were $39,331, or $4,968 (14.5 percent) higher than earnings in the same years of nonmovers ($34,363). Among women aged 45–49 who moved during 2004–2006, postmove earnings were $37,997, or $4,099 (9.7 percent) *lower* than the earnings in the same period among women who did not move ($42,096).

Overall, Charts 9 and 10 show that men and women younger than 40 who moved to a new county or state had higher real earnings 4–6 years after moving than those who did not move, even though their earnings before the move had been lower than those of nonmovers. For men and women aged 40–49, the opposite was true: Those who remained in the same location had higher earnings than movers in the period 4–6 years after the movers relocated to a new county or state. These results suggest that if an advantage in earnings growth accrues to those who move to a new location, it appears to occur mainly among workers younger than 40. One possible explanation for this finding is that people of different ages may move for different reasons. For example, younger people may move mainly in order to find better-paying employment, while older people might be more likely to move to be closer to family members in need of child care or elder care.

## *Multivariate Analysis of Earnings Change*

Charts 7 through 10 show the average earnings of men and women by age in two different periods; however, earnings also vary with other personal characteristics and with local economic conditions. Tables 3 and 4 show, for men and women respectively, the results of ordinary least square (OLS) regressions in which the dependent variable is the change in the logarithm of real 3-year mean annual earnings between two periods, controlling for geographic mobility and other factors. Table 3 shows the results of a regression for men who either moved to a new county or state in one of two observation periods and remained in that location for at least 5 years or did not move during that period and remained in the same location for at least 5 years. The upper panel presents regression results for men born from 1945 through 1971 for the 1994–1996 observation period; the lower panel does so for men born from 1955 through 1981 for the 2004–2006 observation period. Table 4 presents the same parameters for women.

The dependent variable in the model is the change in the natural logarithm of real mean annual earnings between two 3-year periods. The change in the logarithm of earnings is approximately equal to the percentage change in earnings. The first observation period for those who moved is the 3 years up to and including the year of the move; mean earnings for nonmovers are calculated for the same 3-year period. The second observation period for those who moved consists of the fourth, fifth, and sixth years after the move; again, mean earnings for nonmovers are calculated for the same 3-year period.

The model includes the following conditional independent variables:

- Whether the individual moved to a new county or state (=1) or not (=0);
- Whether the individual changed employers (=1) or not (=0);
- Whether the individual both moved and changed employers (=1) or not (=0);
- Whether the individual is non-Hispanic white (=1) or not (=0);
- Whether the individual was born outside the United States (=1) or not (=0); and
- Whether the individual's county of residence (for movers, the former residence) is classified as metropolitan (=1) or not (=0).

The model also includes the following categorical independent variables:

- Quartile rank of the individual's average annual earnings in the 3 years prior to moving (or not moving) for persons of the same sex and 5-year age group. The first (lowest) quartile is the omitted category.
- The individual's birth cohort. The youngest cohort is the omitted category.
- The region of the individual's county of residence (for movers, the former residence). Northeast is the omitted category.

Finally, the model also includes two continuous independent variables. One represents the ratio of the 3-year average unemployment rate in the person's

**Table 3.**
**OLS regressions for change in logarithm of men's 3-year average earnings from event year *n* to *n* +6**

| Independent variable | Parameter estimate | Standard error | *t* value | *p*-value [a] |
|---|---|---|---|---|
| | ***1994–1996 observation period (1945–1971 birth cohorts)*** [b] | | | |
| Intercept | 0.6788* | 0.0117 | 57.90 | <0.0001 |
| Moved to other county or state | 0.0480* | 0.0083 | 5.76 | <0.0001 |
| Changed employer | 0.0481* | 0.0039 | 12.41 | <0.0001 |
| Moved *and* changed employer | 0.0894* | 0.0120 | 7.45 | <0.0001 |
| Quartile of mean annual earnings | | | | |
| 4th (highest) | -0.5288* | 0.0044 | -120.37 | <0.0001 |
| 3rd | -0.5223* | 0.0044 | -117.81 | <0.0001 |
| 2nd | -0.4295* | 0.0045 | -95.79 | <0.0001 |
| Birth cohort | | | | |
| 1945–1949 | -0.4416* | 0.0047 | -93.49 | <0.0001 |
| 1950–1954 | -0.3790* | 0.0046 | -82.22 | <0.0001 |
| 1955–1959 | -0.3225* | 0.0046 | -70.61 | <0.0001 |
| 1960–1964 | -0.2284* | 0.0046 | -49.15 | <0.0001 |
| White, non-Hispanic | 0.0808* | 0.0035 | 23.20 | <0.0001 |
| Foreign place of birth | 0.0984* | 0.0050 | 19.84 | <0.0001 |
| Region | | | | |
| Midwest | -0.0033 | 0.0043 | -0.76 | 0.4473 |
| South | -0.0181* | 0.0042 | -4.26 | <0.0001 |
| West | -0.0028 | 0.0043 | -0.66 | 0.5093 |
| Metropolitan county | 0.0466* | 0.0045 | 10.30 | <0.0001 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | -0.0146* | 0.0043 | -3.43 | 0.0006 |
| Median household income | 0.0997* | 0.0066 | 15.07 | <0.0001 |
| | ***2004–2006 observation period (1955–1981 birth cohorts)*** [c] | | | |
| Intercept | 0.3903* | 0.0136 | 28.64 | <0.0001 |
| Moved to other county or state | 0.0532* | 0.0084 | 6.33 | <0.0001 |
| Changed employer | -0.0040 | 0.0040 | -1.01 | 0.3125 |
| Moved *and* changed employer | 0.1205* | 0.0125 | 9.65 | <0.0001 |
| Quartile of mean annual earnings | | | | |
| 4th (highest) | -0.2910* | 0.0043 | -67.16 | <0.0001 |
| 3rd | -0.2903* | 0.0042 | -69.01 | <0.0001 |
| 2nd | -0.2636* | 0.0041 | -63.77 | <0.0001 |
| Birth cohort | | | | |
| 1955–1959 | -0.3940* | 0.0048 | -81.62 | <0.0001 |
| 1960–1964 | -0.3452* | 0.0048 | -71.57 | <0.0001 |
| 1965–1969 | -0.2861* | 0.0049 | -58.38 | <0.0001 |
| 1970–1974 | -0.1894* | 0.0051 | -37.51 | <0.0001 |
| White, non-Hispanic | 0.0817* | 0.0036 | 22.47 | <0.0001 |
| Foreign place of birth | 0.0905* | 0.0046 | 19.64 | <0.0001 |
| Region | | | | |
| Midwest | -0.0687* | 0.0044 | -15.46 | <0.0001 |
| South | -0.0068 | 0.0042 | -1.60 | 0.1096 |
| West | -0.0018 | 0.0045 | -0.39 | 0.6965 |
| Metropolitan county | 0.0289* | 0.0289 | 10.30 | <0.0001 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | 0.0082 | 0.0067 | 1.22 | 0.2225 |
| Median household income | 0.0959* | 0.0070 | 13.70 | <0.0001 |

SOURCE: Author's calculations using CWHS data.

NOTES: For movers, "region" and "county" refer to prior location.

* = statistically significant at the 0.01 level.

a. Two-tailed test.

b. Sample size = 181,768. Dependent mean = 0.248. $R^2$ = 0.1655.

c. Sample size = 199,618. Dependent mean = 0.109. $R^2$ = 0.0780.

**Table 4.**
**OLS regressions for change in logarithm of women's 3-year average earnings from event year *n* to *n* +6**

| Independent variable | Parameter estimate | Standard error | *t* value | *p*-value [a] |
|---|---|---|---|---|
| | ***1994–1996 observation period (1945–1971 birth cohorts)*** [b] | | | |
| Intercept | 0.8151* | 0.0148 | 55.00 | <0.0001 |
| Moved to other county or state | -0.0223 | 0.0116 | -1.93 | 0.0536 |
| Changed employer | 0.0798* | 0.0048 | 16.74 | <0.0001 |
| Moved *and* changed employer | 0.0805* | 0.0164 | 4.90 | <0.0001 |
| Quartile of mean annual earnings | | | | |
| 4th (highest) | -0.8281* | 0.0051 | -162.80 | <0.0001 |
| 3rd | -0.7771* | 0.0050 | -156.15 | <0.0001 |
| 2nd | -0.6742* | 0.0049 | -136.76 | <0.0001 |
| Birth cohort | | | | |
| 1945–1949 | -0.2911* | 0.0060 | -48.87 | <0.0001 |
| 1950–1954 | -0.2043* | 0.0058 | -35.36 | <0.0001 |
| 1955–1959 | -0.1617* | 0.0058 | -28.07 | <0.0001 |
| 1960–1964 | -0.1479* | 0.0059 | -25.05 | <0.0001 |
| White, non-Hispanic | 0.0078 | 0.0042 | 1.85 | 0.0643 |
| Foreign place of birth | 0.0462* | 0.0065 | 7.10 | <0.0001 |
| Region | | | | |
| Midwest | 0.0089 | 0.0055 | 1.62 | 0.1052 |
| South | -0.0211* | 0.0053 | -3.94 | 0.0001 |
| West | 0.0154* | 0.0055 | 2.82 | 0.0048 |
| Metropolitan county | 0.0621* | 0.0057 | 10.88 | <0.0001 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | 0.0020 | 0.0054 | 0.37 | 0.7114 |
| Median household income | 0.1233* | 0.0084 | 14.76 | <0.0001 |
| | ***2004–2006 observation period (1955–1981 birth cohorts)*** [c] | | | |
| Intercept | 0.5663* | 0.0154 | 36.02 | <0.0001 |
| Moved to other county or state | -0.0001 | 0.0104 | -0.01 | 0.9920 |
| Changed employer | 0.0119* | 0.0044 | 2.71 | 0.0067 |
| Moved *and* changed employer | 0.1064* | 0.0152 | 7.01 | <0.0001 |
| Quartile of mean annual earnings | | | | |
| 4th (highest) | -0.6197* | 0.0047 | -131.67 | <0.0001 |
| 3rd | -0.5861* | 0.0046 | -127.13 | <0.0001 |
| 2nd | -0.5288* | 0.0046 | -115.55 | <0.0001 |
| Birth cohort | | | | |
| 1955–1959 | -0.2558* | 0.0053 | -48.26 | <0.0001 |
| 1960–1964 | -0.1911* | 0.0053 | -35.83 | <0.0001 |
| 1965–1969 | -0.1652* | 0.0055 | -30.13 | <0.0001 |
| 1970–1974 | -0.1458* | 0.0056 | -25.89 | <0.0001 |
| White, non-Hispanic | 0.0581* | 0.0039 | 15.03 | <0.0001 |
| Foreign place of birth | 0.0884* | 0.0052 | 16.98 | <0.0001 |
| Region | | | | |
| Midwest | -0.0526* | 0.0049 | -10.65 | <0.0001 |
| South | -0.0183* | 0.0047 | -3.88 | 0.0001 |
| West | -0.0004 | 0.0051 | -0.07 | 0.9442 |
| Metropolitan county | 0.0432* | 0.0053 | 8.16 | <0.0001 |
| Ratio of origin county to national— | | | | |
| Unemployment rate | 0.0094 | 0.0075 | 1.26 | 0.2077 |
| Median household income | 0.1221* | 0.0079 | 15.47 | <0.0001 |

SOURCE: Author's calculations using CWHS data.

NOTES: For movers, "region" and "county" refer to prior location.

* = statistically significant at the 0.01 level.

a. Two-tailed test.

b. Sample size = 182,138. Dependent mean = 0.297. $R^2$ = 0.1884.

c. Sample size = 204,976. Dependent mean = 0.177. $R^2$ = 0.1211.

county of residence (for movers, the former residence) to the national 3-year average unemployment rate. The other represents the ratio of median household income in the person's county of residence (for movers, the former residence) to national median household income.

Table 3 shows that, other things being equal, moving to a new county or state had a small but statistically significant positive relationship with the change in the logarithm of men's 3-year mean earnings for both the 1994–1996 and the 2004–2006 movers. The coefficient for the *moved* variable increased slightly between the two periods. The coefficient for the *changed employer* variable was positive and significant for 1994–1996 movers. For 2004–2006, the coefficient was negative but not significant. The geographic-mobility and employer-change interaction variable was positive and significant in both periods. These results do not support the hypothesis that diminishing earnings gains from moving contributed to declining geographic mobility during this period.

The regression results for both observation periods also indicate that, all else being equal, men in the lowest earnings quartile experienced greater percentage gains in earnings than men in higher earnings quartiles, and men in the youngest age group experienced greater percentage gains in earnings than older men. In both periods, non-Hispanic white men experienced greater percentage gains in earnings than men in other racial/ethnic groups, and foreign-born men experienced greater percentage gains in earnings than native-born workers.

In both periods, men who resided in metropolitan counties experienced larger percentage increases in earnings than those in nonmetropolitan counties. The coefficient for high county unemployment relative to the national rate was negative and statistically significant in 1994–1996 but was not significant in 2004–2006. The coefficient for high county median household income relative to national household income was positive and statistically significant in both periods. In other words, men who lived in counties with above-average median household income experienced greater percentage increases in earnings than those who lived in lower-income counties, other things being equal.

Table 4 shows the regression results for women, which differ from those for men in an important respect: The independent *moved* variable was negative for both 1994–1996 and 2004–2006 movers, but the coefficient was not statistically significant in either period. Both the variable indicating a change in employer and the variable interacting geographic mobility *and* employer change were positive and statistically significant. These results suggest that, for women, employer change alone and geographic mobility combined with employer change were positively correlated with earnings gains, but geographic mobility alone was not.

Similar to men, in both periods, women in the lowest earnings quartile experienced greater percentage gains in earnings than workers in higher earnings quartiles, all else being equal. Likewise, women in the youngest age group experienced greater percentage gains in earnings than older women. The change in earnings for non-Hispanic white women was not significant for the 1994–1996 period, but was positive and significant for the 2004–2006 period. In both periods, foreign-born women experienced greater percentage gains in earnings than native-born women. As was also the case with men, in both periods, women in metropolitan counties experienced greater percentage increases in earnings than women in nonmetropolitan counties.

Earnings changes for women in counties with higher unemployment rates than the national rate were not statistically significant during either period. Earnings changes for women in counties with higher median household income than national median household income were positive and significant in both periods, as they were for men. Thus, women who lived in counties with above-average median household income experienced greater percentage increases in earnings than those who lived in lower-income counties, other things being equal.

## Summary and Conclusion

This article uses CWHS data to examine trends in geographic mobility and employer change in the United States and to compare the annual earnings of movers and nonmovers over time. The data show that the average annual percentage of men and women aged 25–49 who moved to a new county or state declined substantially between 1994–1996 and 2014–2016. The decline occurred among both younger and older workers, but was larger among men and women younger than 40. The majority of the decline in geographic mobility rates among men and women occurred between 2004–2006 and 2014–2016. In contrast with the decline in annual rates of geographic mobility, there was little change in the average annual percentage of workers who changed employers during that span. Among men, average annual rates of

employer change were relatively stable, while among women they rose slightly. This result contrasts with the findings of some studies, which have reported a downward trend in employer change by American workers.

Among workers who moved to a new county or state, a majority also changed employers, and the proportion of movers who also changed employers was relatively stable between 1994–1996 and 2014–2016. By contrast, among men and women who changed employers, the percentages who also moved to another county or state declined substantially between 1994–1996 and 2014–2016.

Multivariate analysis indicates that younger workers, those with recent 3-year mean earnings in the lowest earnings quartile for their 5-year age group, and those who resided in counties with above-average median household income were relatively more likely to have moved. This was true both for the full samples of men and women and for the subsamples of those who had changed employers in the previous year. Among men and women who had *not* recently changed employers, those in higher earnings quartiles were slightly more likely to have moved than were those in the lowest quartile, possibly because they had been transferred or moved to another county within the same commuting area while remaining with the same employer.

In both 1994–1996 and 2004–2006, among men aged 25–49 who moved, median 3-year mean annual earnings before moving were lower than earnings in the same period among men who did not move. Among women who moved, median 3-year mean annual earnings before moving were lower than those of nonmovers over the same period in all five age groups in 1994–1996 and in three of five age groups during 2004–2006.

For men, gains in earnings after moving occurred mainly among those younger than 40. For those who moved in 1994–1996 or in 2004–2006, real annual earnings 4–6 years after moving were higher than those of men in the same age group who did not move, even though their premove median 3-year mean earnings were lower than those of nonmovers. By contrast, in both periods, men aged 40–49 who moved had *lower* real annual earnings 4–6 years after moving than men who did not move. Among women, too, gains in earnings after moving appear to have occurred mainly among those younger than 40, while women aged 40–49 who moved had lower earnings 4–6 years later than similarly aged women who did not move.

An OLS regression on the change in the logarithm of 3-year real mean annual earnings over time shows that for men, moving to a new county or state in either 1994–1996 or 2004–2006 had a small but statistically significant positive relationship with the change in earnings, other things being equal. The coefficient for the *moved* variable increased slightly between the two periods, suggesting that the gain in earnings associated with geographic mobility increased during that span. The coefficient on the *changed employer* variable was positive and significant in 1994–1996 but not in 2004–2006. The coefficient for the variable interacting both geographic mobility and employer change was positive and significant in both periods. Overall, the results do not support the hypothesis that diminishing earnings gains from moving contributed to declining geographic mobility of men in the United States during this time. In the regression on the change in women's earnings, the independent *moved* variable was not statistically significant in either 1994–1996 or 2004–2006. The *changed employer* variable was positive and significant in both periods, as was the variable interacting geographic mobility and employer change.

In summary, data from the CWHS reveal that the annual average proportion of men and women aged 25–49 who moved to a new county or state declined from 1994–1996 through 2014–2016, while the annual average proportion who changed employers remained relatively stable. Among men and women younger than 40, those who moved in 1994–1996 or 2004–2006 had higher 3-year average earnings 4–6 years later than those who did not move. Among men, moving to a new county or state was positively and significantly correlated with higher earnings 4–6 years later. Among women, the relationship between moving and earnings was not statistically significant in either period. The results suggest that the decrease in geographic mobility rates during this period is unlikely to have been caused by declining gains in annual earnings among those who moved.

## Notes

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[1] As an example of such a secondary effect, Karahan and Rhee estimated that when the share of workers aged 40 to 60 in a state increases, that age group's lower migration rate tends to lower the migration rate of all workers in the state because firms recruit primarily from the local labor market.

[2] The Bureau of Labor Statistics (2018) defines displaced workers as "persons 20 years of age and older who lost or left jobs because their plant or company closed or moved, there was insufficient work for them to do, or their position or shift was abolished."

[3] The CWHS comprises two components, known as the active file and the inactive file. The active file contains the earnings records for workers with earnings from any employment (including self-employment), regardless of whether those earnings were covered under Social Security. The inactive file contains records only for workers who never had covered earnings posted to the Master Earnings File.

[4] The Department of Agriculture's Economic Research Service (2019) defines commuting zones as geographic units that reflect the local economy where people live and work.

[5] CWHS data for years before 1993 lack a variable that permits the researcher to identify whether the geographic code indicates the employee's place of residence or the employer's location.

[6] The Internal Revenue Service (IRS) issues an EIN for an individual firm, whether it is organized as a corporation, partnership, or sole proprietorship. If a firm has establishments in multiple locations, all such establishments have the same EIN. In some cases, the IRS may issue a new EIN for a given firm. Because these instances represent a small percentage of EIN changes in any given year, they do not greatly distort the estimated incidence of workers changing employers.

[7] Of the five demographic traits most commonly used as regressors—age, sex, race, marital status, and education—the CWHS includes variables describing only the first three.

[8] The observations represent person-years observed for a given subset of the 3,467,451 person-records in the CWHS data file.

[9] Separate analysis showed a correlation coefficient of 0.81 between state median household income and the proportion of adult state residents that had earned a bachelor's or higher degree.

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# The Benefit Receipt Patterns and Labor Market Experiences of Older Workers Who Were Denied Social Security Disability Insurance Benefits on the Basis of Work Capacity

by Jody Schimmel Hyde, April Yanyuan Wu, and Lakhpreet Gill*

*This article examines the experiences of Social Security Disability Insurance (DI) applicants aged 51 or older who were initially denied benefits because the disability examiner determined that they could perform either their past work or other work. We use Health and Retirement Study survey data linked to administrative data on benefit application and receipt and earnings from the Social Security Administration. We find that few older DI applicants who were denied benefits on this basis resumed work at a substantial level following denial. More commonly, applicants denied at this stage continued to pursue benefits, often successfully. Nearly two-thirds of initial work capacity-related denials were ultimately allowed DI benefits after appealing the initial decision or reapplying, and our estimates suggest that many of the rest claimed Old-Age and Survivors Insurance benefits before they reached full retirement age.*

## Introduction

As workers approach retirement, they are more likely to experience a health condition that could limit their ability to remain employed. Among individuals aged 51 to 55 in 1992, one-quarter reported experiencing a health condition that, by age 62, had limited their ability to work at least once (Johnson, Mermin, and Murphy 2007). A worker becomes eligible for unreduced Old-Age and Survivors Insurance (OASI) benefits, commonly known as Social Security retirement benefits, on reaching his or her full retirement age (FRA: 65 to 67, depending on year of birth). If a new health condition significantly affects a worker's ability to remain in the labor force, and the worker has not yet reached FRA, he or she may be eligible for Social Security Disability Insurance (DI) benefits. Alternatively, a worker who has reached age 62 (but not FRA) may decide to claim actuarially reduced OASI benefits. Early (pre-FRA) OASI claiming reduces a monthly benefit by as much as 30 percent (if claimed at age 62 by a worker whose FRA is 67).[1] When workers claim early OASI retirement benefits, the Social Security Administration (SSA) asks them whether they have experienced a health condition, injury, or illness in the past 14 months that left them unable to work. When applicants report any such experience, SSA considers their eligibility for DI benefits. Doing so is relatively costless for those who have already stopped working, and if DI benefits are awarded, the applicant will receive a higher monthly retirement benefit upon

**Selected Abbreviations**

| | |
|---|---|
| CPI | Consumer Price Index |
| CYBF | Cross-Year Benefits file |
| DDS | Disability Determination Service |
| DI | Disability Insurance |
| FRA | full retirement age |
| HRS | Health and Retirement Study |

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**Selected Abbreviations—*Continued***

| | |
|---|---|
| O*NET | Occupational Information Network |
| OASI | Old-Age and Survivors Insurance |
| OIG | Office of the Inspector General |
| PRW | past relevant work |
| RFC | residual functional capacity |
| SEF | Summary Earnings file |
| SGA | substantial gainful activity |
| SSA | Social Security Administration |
| SSI | Supplemental Security Income |

reaching FRA than he or she would receive with an early-claiming reduction. However, an award of disability benefits is far from certain.

This study examines the employment, earnings, and benefit-receipt outcomes of workers aged 51 or older whose applications for DI disabled-worker benefits are denied by the SSA Disability Determination Service (DDS) in the initial review levels.[2] We find that almost half of the applicants aged 51 or older are initially denied benefits. In addition to documenting the employment, earnings, and benefits trajectory of these applicants, we compare outcomes by reason for denial.

SSA uses a five-step sequential evaluation process (detailed later) to determine whether an applicant meets the criteria for benefit award. In the first three steps, evaluators assess the applicant's insured status and the medical factors that affect the ability to continue or resume work. We focus on applications denied at the fourth and fifth steps, in which evaluators assess the applicant's work capacity.[3] In most cases, the DDS assesses these applicants to determine whether, in light of their alleged medical impairments, their residual functional capacity (RFC) allows them to perform either past relevant work (PRW), at step 4, or other work, at step 5.[4] In considering RFC at step 5, disability examiners also account for vocational factors—age, education, and work experience—using guidelines known as the "medical-vocational grids."

Our study sheds light on the extent to which DI applicants denied for work-capacity reasons return to work, and compares their employment and earnings trajectories with those of applicants denied for other reasons or whose claims are ultimately allowed. Strictly speaking, the DI *applications,* and not the applicants themselves, are allowed or denied in the disability determinations. However, for practical reasons, we use "denied applicants" as shorthand to refer to our study sample, whose members applied for DI disabled-worker benefits, had their applications denied by the DDS at the initial review level, and were older than 50 and younger than FRA at the time of the initial denial.

We find that relatively few older denied applicants return to work, and highlight reasons why they might not. In particular, we examine the shares of initially denied applicants that appeal the denial or reapply for DI, and the shares that are ultimately awarded benefits. We also identify how many denied applicants claim OASI benefits after reaching the earliest age of eligibility (62) but prior to FRA—an option not available to younger workers, but a particularly salient one for older workers. We find some evidence that employment paths differ based on the sequential-evaluation step at which the initial denial occurred, and that the likelihood of working declines as time passes after the decision for all older denied applicants. Our findings also point to the possibility that applicants denied at step 5 who returned to work had slightly higher postdenial earnings than did those who were denied at other steps.

To conduct our analysis, we used Health and Retirement Study (HRS) data linked to SSA records on earnings and benefit application and receipt. Our linked sample consists of 805 HRS respondents who applied for DI at age 51 or older but younger than FRA from 1992 to 2012; of these, only 384 applicants were denied benefits, with 259 denied for work-capacity reasons. Although our sample is admittedly small, the richness of the longitudinal survey data in the HRS offers advantages that would not be possible using administrative data alone, including information about applicant demographics, socioeconomic status, and occupational characteristics. Using information from agency administrative records, we stratified DI applicants based on the outcome of and the basis for the initial decision. We measured the employment and earnings of groups of denied applicants using the HRS Respondent Cross-Year Summary Earnings file (hereafter, the Summary Earnings file, or SEF), which links survey results with earnings data from SSA's Master Earnings File. Finally, to better understand postdenial employment patterns, we used data from the linked SSA records on applicants' subsequent receipt of DI and OASI benefits.

Before describing our study, we outline the SSA disability determination process and review existing evidence on the earnings of denied applicants. We then describe our approach to identifying DI applicants using the linked HRS data and administrative records

and examine the characteristics of denied applicants with different reasons for initial denial. We go on to consider applicants' benefit trajectories after denial, and follow that by tracing their employment and earnings trajectories both before application and after adjudication. We then conclude and discuss implications for policy.

## The Disability Determination Process

An individual is eligible for DI benefits if he or she is unable to engage in substantial gainful activity (SGA) because of a medically determinable physical or mental impairment that has lasted or is expected to last for a continuous period of at least 12 months or is expected to result in death (SSA 2012b). Earnings that meet or exceed an annually adjusted monthly earnings threshold signify SGA. In 2020, SGA is indicated by $1,260 for nonblind individuals and $2,110 for blind individuals. SSA disability examiners follow the five-step sequential evaluation process outlined below to determine eligibility.

In step 1, the examiners determine whether the applicant's work and payroll-tax contribution history is sufficient to qualify him or her as insured. If so, disability examiners assess whether the applicant has an impairment that meets the medical eligibility criteria in step 2. If the medical evidence does not establish that the applicant's conditions are severe or will last for at least 12 months, the examiner denies the claim for medical reasons. In this article, we refer to these as "medical denials."

If an impairment is deemed to be sufficiently severe in step 2, the examiner goes on to determine whether it satisfies the medical criteria for specific conditions contained in SSA's Listing of Impairments in step 3. The listing includes hundreds of conditions that result in a benefit award if sufficiently severe. If an applicant has more than one documented impairment, the examiner may find that the impairments in combination are equivalent to the medical criteria in a given listing. The adjudicator allows the application if the impairments are determined to meet or "equal" the listings.

If an application is not allowed at step 3, the examiner assesses the applicant's RFC (for details, see SSA 2018b) and compares it to the requirements of the applicant's PRW (see SSA 2011) at step 4.[5] If the examiner deems that the applicant is able to perform PRW (which extends to any work performed within the last 15 years), the application is denied at step 4. Examiners compare RFC with PRW on a function-by-function basis; in other words, they compare the requirements of the past work, using information provided by the applicant or contained in SSA's Dictionary of Occupational Titles, with the applicant's RFC (for details, see SSA 2017b and SSA 2017c). Examiners consider the applicant's ability to perform PRW both as it was performed in his or her job (step 4a) or as it is generally performed in the national economy (step 4b). Assessing the ability to perform PRW does not account for the availability of such work in the current economy or for other economic conditions that might affect applicants' ability to find work (see SSA 2011).

The fifth step affects applicants whose RFCs are deemed incompatible with the performance of PRW. The cases of nearly half of insured DI applicants reach this final step (Wixon and Strand 2013; Mann, Stapleton, and de Richemond 2014). The examiner assesses ability to perform other work by comparing the applicant's RFC to the exertional requirements of work outlined in SSA's medical-vocational guidelines (see SSA 2015). Examiners deny benefits for applicants whose RFCs indicate that they can meet the requirements. As in step-4 decisions, step-5 medical determinations do not account for current economic conditions.

By law, SSA must consider the vocational factors—age, education, and work experience—in determining whether an applicant can engage in substantial work, although the law does not specify how those factors should be incorporated into the determination process. The grids show the type of work that can be done given the applicant's RFC and vocational factors. The types of work are grouped by exertional requirements, including sedentary, light, medium, heavy, or very heavy. Because the grids vary by age, with separate factors for ages 50–54, 55–59, and 60–64, they are particularly salient for older workers. Benefit awards are more likely with age because of the grids, even holding disability, education, and work experience constant. For example, a 54-year-old individual who could perform sedentary work would not receive an award based on the grid for 50- to 54-year-olds but could receive an allowance a year later when the examiner consults the grid for 55- to 59-year-olds, under which the individual is deemed unable to work. One final important factor in a step-5 determination is whether the applicant has transferable skills from PRW that enable him or her to meet the requirements of other occupations available in the RFC category. If the applicant is deemed to have transferable skills, that factor dominates age and education in reaching an award determination (Johns n.d.).

SSA's use of vocational factors and its assessment of ability to perform PRW have been criticized for many years. In reviewing the relationship between vocational factors and employment, Mann, Stapleton, and de Richemond (2014) did not find evidence in the literature that vocational factors alone can predict ability to perform work that one has not performed before—independent of other factors that might predict future work. For that reason and others, vocational factors have drawn policymakers' attention and reform proposals (for example, Warshawsky and Marchand 2015). SSA has also grappled with identifying the demands of work when comparing an applicant's RFC and the requirements of PRW. The agency continues to use the Dictionary of Occupational Titles, which is outdated and often does not reflect current job demands. More recently, SSA sponsored a new data set, the Occupational Information System (OIS), which is based on occupational information collected by the Bureau of Labor Statistics and is updated every 5 years.[6]

## *Prior Evidence About the Earnings of Denied Applicants*

The earnings of denied DI applicants were the subject of a robust literature nearly three decades ago (for example, Bound 1989 and Parsons 1991). More recently, von Wachter, Song, and Manchester (2011) reexamined the topic, taking advantage of improvements in data and processing. However, those studies documented the earnings of denied applicants to benchmark what might have been expected for the earnings of allowed applicants in the absence of benefits. As such, the studies aggregated denied applicants and did not explore employment and earnings variations by reason for denial.

Two other recent studies focused on applicants denied because of their work capacity, which therefore relate closely to our analysis. However, the timeframe and applicant selection criteria used in those studies differed from ours. In one study (SSA 2017a), the agency's Office of the Inspector General (OIG) determined that less than half of applicants denied at steps 4 and 5 in 2013 returned to work. Moreover, average earnings after denial (among those with earnings) were less than $9,400 per year, or about 70 percent of that group's $13,640 preapplication average annual earnings. The OIG findings align with those of an earlier study (Strand and Trenkamp 2015), which focused on step-5 denials, although the studies analyzed different populations and observation periods.[7] Strand and Trenkamp determined that employment declined by 22 percentage points (26 percent) from 2000 to 2008 among those who were initially denied in 2005. The authors also documented important differences in outcomes by preapplication earnings decile, but they found that postdenial earnings (both median and maximum) were persistently lower than preapplication earnings in nearly every decile, averaging around 77 percent for the sample overall.

Our work complements and extends the earlier studies along two important dimensions. First, we separate step-4 denials from step-5 denials and compare outcomes between those groups as well as with other groups of denied applicants. By contrast, the 2017 OIG study combined step-4 and step-5 denials, while Strand and Trenkamp considered only step-5 denials; in addition, neither study compared outcomes with those of other applicant groups. Second, we link administrative data to longitudinal survey data, allowing us to measure attributes such as health, household, and occupational characteristics—attributes that neither of the earlier studies measured because they are not contained in administrative data. This information provides a richer picture of denied applicants and how they differ from their allowed counterparts.

Although it has many advantages, the HRS limits our consideration of applicants to those aged 51 or older, whereas earlier studies considered applicants of all ages. As described earlier, vocational factors at step 5 are progressively less stringent for applicants as they age. As a result, the reasons we find for denials based on work capacity split differently from those found in other studies, reflecting the consideration of age in the vocational grid. For example, of denials based on work capacity in our study, more than two-thirds occurred at step 4 (shown later), in contrast with the OIG report, which found that only one-quarter of denials for work capacity occurred at step 4 and three-quarters occurred at step 5. Because the OIG report considered Supplemental Security Income (SSI) applicants as well as those for DI (and concurrent SSI and DI) benefits, its findings differ from ours, as postdenial employment for older DI applicants with substantial work histories will likely differ from those of SSI applicants of all ages without such experience. Given that our sample does not include younger applicants, our findings are not directly comparable with those of the earlier studies.

## Data and Sample Selection

Our study capitalizes on a significant linkage of survey and administrative data. We use the HRS, a nationally representative longitudinal survey of Americans collected by the Institute for Social Research at the University of Michigan. The HRS interviews respondents aged 51 to 61 then conducts biennial follow-up interviews with them on a range of subjects. We use data collected in the 1992 through 2012 waves of the HRS. We link the HRS survey responses to administrative data from several SSA sources, with which we measure DI applications and benefit receipt and collect annual earnings data through 2012.[8] From the Department of Labor's Occupational Information Network (O*NET), we also collect data on occupational characteristics.[9] By linking the survey results with the data from various administrative sources, we are able to compile detailed information and avoid the reporting errors that are common with self-reported benefit receipt and earnings.

Our estimates are weighted to account for both the complex survey design and the respondents' consent to having their survey responses linked to administrative data. We use the weights provided by the HRS.[10,11] The HRS consent weights account for nonrandom selection into the group that consented to the administrative data linkage. Even though more than three-quarters of the respondents in our observation period agreed to the linkage,[12] evidence has shown that consenters differed from the full HRS sample in terms of age, race, sex, income, and education (see, for example, Gustman and Steinmeier 2001; Haider and Solon 2000; and Kapteyn and others 2006).

### Identifying DI Applicants Using the Linked HRS-SSA Files

Our analysis focuses on 805 HRS respondents who applied for DI disabled-worker benefits (that is, benefits based on their own work history) at least once after age 50 and before FRA, following their first HRS interview.[13] The DI applications were relatively evenly distributed from the first observation year (1992) to the last (2012). About one in four (26 percent) of these applicants concurrently applied for SSI. We identified these applicants using SSA's 831 file (named for the Disability Determination and Transmittal form SSA-831), which records information about all applications that receive a medical determination. We used the date of initial application recorded in that file, which we then aligned with the timing of the survey interviews in the linked HRS data to identify the date of first application that followed the first HRS interview.

Importantly, a denial recorded in the 831 file does not mean that the applicant was not ultimately awarded benefits, because the file does not include information about appeals to an administrative law judge (ALJ) or a federal court. Thus, we have information only on the outcomes adjudicated by the DDS. If an application was initially denied but was then reconsidered by the DDS at the request of the applicant, the result of the reconsideration is in the 831 file. In such cases, we used the reconsideration outcome—and the basis for that outcome—to classify the case. To track postdenial benefit trajectories, we developed an approach to identify subsequent ALJ benefit awards using the HRS Cross-Year Benefits file (CYBF), discussed later.

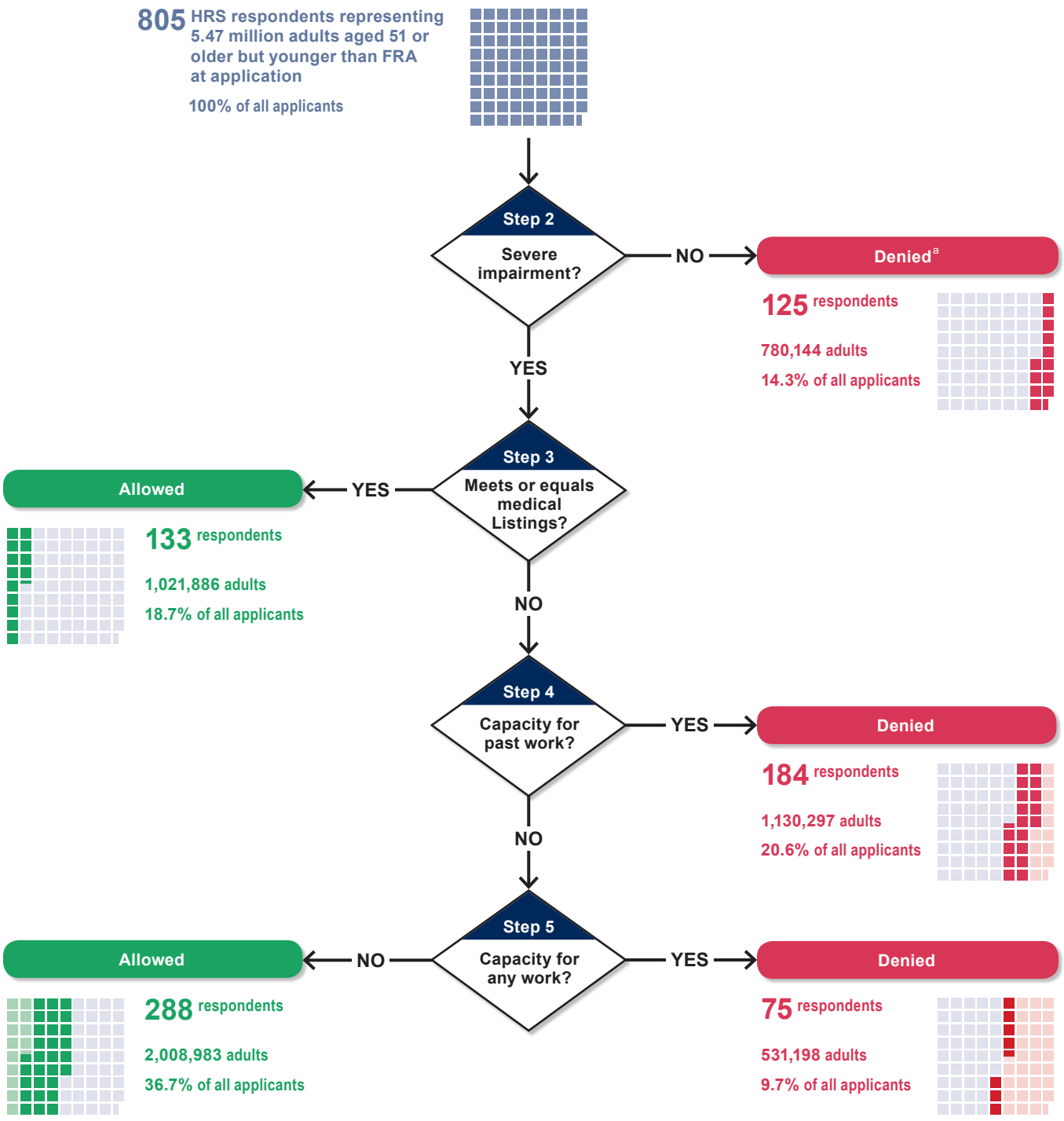
The fact that a respondent's first HRS interview is conducted no earlier than age 51 and as late as age 61 has two methodological implications for our analysis. First, the applications we observe do not necessarily reflect individuals' first applications for DI benefits; of the 805 applicants in our sample, approximately 10 percent had applied for DI at least once from 1988 (the earliest year for which SSA-831 data are available) to the time of their first HRS interview. First interviews in our sample occurred between 1992 (the first HRS year) and 2012 (the last year for which many SSA data are available, although the earnings data continue through 2013).[14] Second, because the age at the first HRS interview ranges from 51 to 61, the first time we observe a DI applicant is left-censored. For instance, if a person participated in an HRS interview for the first time at age 56, but had previously applied for DI at age 52, we would only count applications that occurred at age 56 or older. Yet if another respondent had applied at age 52 but was first interviewed for the HRS at age 51, we would include that application in our analysis.

### Categorizing DI Applicants Based on the Outcome of the Initial Determination

We group applicants based on the outcome of their initial decision as recorded in the 831 file. Chart 1 diagrams our sample based on the five-step sequential evaluation process described earlier. More than half of the older DI applicants in our sample were initially allowed benefits, a substantially higher allowance rate than that for DI applicants of all ages (just above 30 percent) found by Wixon and Strand (2013).[15]

In addition to indicating whether an application was initially allowed or denied, the 831 file provides

**Chart 1.**
**Distribution of DI applicants aged 51 or older but younger than FRA, by outcome of the initial determination (including DDS reconsideration), 1992–2012**




SOURCES: Authors' calculations using HRS data linked to SSA's 831 file; Wixon and Strand (2013).

NOTE: Numbers are weighted to account for the complex HRS survey design and the varying probabilities of respondent consent to matching the HRS data to administrative data.

a. Consists primarily of applicants denied because the impairment was deemed not to be severe or likely to last 12 months or result in death. Also includes a small number of applicants (fewer than 10) who did not meet eligibility requirements before step 3 for other reasons, such as failing to follow prescribed treatment, submit to a consultative examination, or provide sufficient supporting evidence.

information on the step at which the examiner decided the case. We use that information to stratify applicants based on the scheme outlined in Wixon and Strand (2013). Among allowed applicants, about one-third were allowed benefits at step 3 because their impairment met or equaled medical criteria in the Listing of Impairments, and the other two-thirds were allowed at step 5 for reasons related to reduced work capacity. Among denied applicants, about one-third were denied at step 2 because their impairment was deemed not to be severe and the remaining two-thirds were denied for work-capacity reasons. Among the latter, denials for the ability to perform past work (step 4) were more than twice as common as denials for the ability to perform other work (step 5). That result is consistent with the sense that if there is cause to deny an application, examiners will aim to identify it at the first applicable step, and with the fact that some cases can be denied at step 4 but not at step 5.

### *Measuring Employment and Earnings Around the Initial DI Decision*

We measure earnings and employment (indicated by having positive earnings) using the linked SEF, available with permission from the HRS. The file includes Master Earnings File information on annual earnings subject to OASI and DI payroll taxes, up to the taxable maximum in each year, as reported to the Internal Revenue Service.[16] We consider earnings in the years before and after application using the application filing and decision dates. Virtually all HRS respondents who consent to the linkage to administrative data from SSA also have a record in the SEF, although not in every year (perhaps reflecting years with no reportable earnings). To align the administrative data on earnings with self-reported household income in the HRS, we follow the HRS convention of using the calendar year preceding the interview date. Using the Consumer Price Index (CPI), we adjust all income and earnings measures to 2012 dollars.

## *Tracing Preapplication and Postdenial Employment, Earnings, and Benefit Patterns*

We open this section by outlining our methodology for tracing the pattern of benefit receipt for DI applicants in the period between the initial denial and FRA. We then describe our approach to measuring the employment and earnings trajectories of denied applicants in the 5 years before they applied for benefits and the 5 years after their initial denial.

We note that data enabling us to observe the benefits trajectory through FRA, and the employment and earnings of applicants for 5 years after denial, were not available for all applicants in our sample. There are two primary reasons for this right-censoring. First, some denied applicants died: 2 percent died within 5 years of the initial decision, and 4 percent died before they reached FRA. Second, some applicants' cases were not decided until shortly before our observation period ended. Applicants who were denied in or close to 2012 (the last year for which we obtained benefit data) had little time to reapply or appeal their decision and qualify for benefits within our observation period. Likewise, applicants who were denied within 5 years of 2013 (the last year for which we obtained earnings data) could not be observed for the full follow-up period. We discuss the magnitude and effects of censoring caused by partial data availability below.

### *Tracing Postdenial Benefit Patterns*

To compile descriptive statistics on postdenial DI benefit trajectories, we link data from SSA's 831 file to information contained in the CYBF. The CYBF includes administrative data from SSA's Master Beneficiary Record and Payment History Update System, both of which provide monthly information on DI and OASI benefit receipt.

The CYBF link was necessary to determine whether applicants appealed an initial denial because the 831 file does not contain information on allowances and denials at higher levels of appeal. We assume that an applicant was granted benefits on appeal if the CYBF record indicates DI benefit receipt after an initial denial and we observe no later DI applications in the 831 file. We assume that an applicant reapplied and was awarded benefits if we observe DI benefit receipt in the CYBF and a subsequent DI application in the 831 file, although we do not know the adjudication level at which the case was allowed. For those whose records indicate a subsequent DI application but no benefit receipt, we assume that applicants reapplied and were denied, although they may have received an allowance after 2012, the last year in the CYBF. We categorize one final group: those who never applied again. This group comprises those with no subsequent application indicated in their 831 file and no DI benefits in the CYBF. It includes applicants who unsuccessfully appealed their initial denial as well as those who did not appeal; the data do not allow us to distinguish between the two groups.

One limitation of the CYBF data inhibits our ability to categorize postdenial DI benefit receipt correctly. The CYBF variable that indicates type of benefit (disability or retirement) has only accounted for conversions from the former to the latter—and enabled historical tracking of such changes—since May 2009. Before then, the prior status was overwritten when the status changed.[17] To the extent that the absence of prior-status information in those records introduces errors in our results, it would lead us to underestimate the share of denied applicants who received DI benefits before reaching FRA. Specifically, we cannot determine whether DI applicants whose claims were initially denied, who became Social Security beneficiaries between age 62 and FRA, and who attained FRA before May 2009 received OASI benefits only or were first awarded DI benefits. Hence, an initially denied applicant who attained FRA before May 2009 and whom we have classified as entering the OASI beneficiary rolls before FRA may have actually entered the DI rolls first, albeit after age 62.

### *Tracing the Employment and Earnings of Denied Applicants Around the Application and Initial Denial*

We estimate average employment rates and earnings for denied applicants from 5 years before application to 5 years after the initial denial was received. Right-censoring because of a lack of available data after the initial decision (because of death or a decision closing within 5 years of 2013, the last year of available earnings data) affects the share of records for which we can observe employment outcomes for the full period. The number of complete records for applicants through the end of the 5th year after denial is 15–25 percent lower than the number of denied applicants in our sample, with missing data most common among medical denials and least common among step-5 denials.

We did not find that censoring led to any systemic differences in average employment rates or earnings levels for denied applicants. We considered the effects of censoring by comparing mean earnings among all denied applicants in their postdenial years to those of denied applicants with an initial decision in 2008 or earlier; for the latter, we had 5 years of postdecision data available (excluding the few cases of applicants who died). Again, using unadjusted values, we did not find that the employment rate of denied applicants varied in any meaningful way when we imposed this restriction. To the extent that average earnings differed for a given group of denied applicants, the direction was not consistent and the magnitude was small relative to the overall standard deviation of earnings. To maintain our already small sample size, our findings are for all initially denied applicants.

## *Characteristics and Preapplication Occupational Attributes of Denied DI Applicants*

The mean age at application across our entire sample was slightly younger than 58; ages ranged from 51 to 65. Most of the applicants we observe sought DI benefits before they were eligible to claim OASI benefits; only 12 percent had reached age 62 and were eligible for OASI when they applied for DI.

Table 1 shows that applicants denied because of work capacity were younger on average (57.4) than those denied for medical reasons (58.2). Consistent with the medical-vocational grids, work-capacity denials were concentrated in higher educational-attainment groups than medical denials; for example, 19 percent of work-capacity denials had less than a high school education, compared with 27 percent of medical denials. Relative to medical denials, work-capacity denials were less likely to be married (58 percent versus 75 percent) or Hispanic (8 percent versus 21 percent). They also had longer tenure in the job they held 2 years before application (10.3 years versus 6.5 years), more years with positive earnings at ages 22 to 50 (21 versus 20), higher average annual earnings in those years ($27,696 versus $19,638), and higher household income ($56,266 versus $45,652). Table 2 shows that work-capacity denials were also more likely to work in occupations requiring computer use (23 percent versus 15 percent), the ability to withstand stress (28 percent versus 20 percent) and to deal with unpleasant or angry people (14 percent versus 3 percent), and basic skills (59 percent versus 49 percent). However, they were less likely to work in jobs requiring social skills (21 percent versus 44 percent) or "system" skills (6 percent versus 17 percent).[18]

Table 1 shows that applicants denied at step 4 were older than those denied at step 5 (on average, 58.2 versus 55.7).[19] Relative to applicants whose initial determination was a denial at step 5, those who were denied at step 4 were more likely to be women (60 percent versus 41 percent), to be Hispanic (10 percent versus 1 percent), or to have a high school education (48 percent versus 23 percent). They were less likely to be married (54 percent versus 66 percent), to be white (69 percent versus 83 percent), or to have any

**Table 1.**
**Demographic, employment, and income characteristics of older DI applicants denied at the initial level, by reason for denial, 1992–2012**

| Characteristic | Denied for medical reasons (step 2) | | Denied for work capacity | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | | Total (step 4 or 5) | | Able to perform past work (step 4) | | Able to perform other work (step 5) | |
| | Statistic | Standard error | Statistic [a] | Standard error | Statistic | Standard error | Statistic [b] | Standard error |
| Number (weighted) | 780,144 | | 1,661,495 | | 1,130,297 | | 531,198 | |
| Demographics | | | | | | | | |
| Average age | 58.2 | 0.3 | 57.4** | 0.2 | 58.2 | 0.2 | 55.7*** | 0.4 |
| Percentage— | | | | | | | | |
| Women | 58.1 | 4.5 | 53.6 | 3.2 | 59.7 | 3.7 | 40.7*** | 5.9 |
| Married | 74.7 | 4.0 | 58.2*** | 3.2 | 54.3 | 3.8 | 66.1* | 5.7 |
| White | 68.6 | 4.2 | 73.6 | 2.8 | 69.1 | 3.4 | 83.0** | 4.5 |
| Black | 18.0 | 3.5 | 20.2 | 2.5 | 22.5 | 3.1 | 15.1 | 4.3 |
| Hispanic | 20.5 | 3.7 | 7.5*** | 1.7 | 10.4 | 2.3 | 1.2** | 1.3 |
| Percentage with— | | | | | | | | |
| Less than high school diploma | 27.1 | 4.0 | 18.8* | 2.5 | 17.9 | 2.9 | 20.7 | 4.9 |
| High school diploma | 39.2 | 4.4 | 39.6 | 3.1 | 47.7 | 3.7 | 22.5*** | 5.0 |
| More than high school diploma | 33.7 | 4.3 | 41.6 | 3.1 | 34.4 | 3.5 | 56.8*** | 6.0 |
| Employment status and work experience | | | | | | | | |
| Percentage— | | | | | | | | |
| Employed | 62.5 | 4.6 | 65.8 | 3.1 | 66.2 | 3.7 | 64.9 | 6.0 |
| Employed full time | 71.8 | 5.4 | 73.5 | 3.5 | 74.8 | 4.1 | 70.6 | 7.0 |
| Self-employed | 13.2 | 4.1 | 12.3 | 2.6 | 10.3 | 2.8 | 16.9 | 5.8 |
| Working for a firm with— | | | | | | | | |
| More than 500 employees | 11.8 | 4.2 | 18.5 | 3.5 | 25.9 | 4.8 | 4.6*** | 3.4 |
| Fewer than 100 employees | 67.8 | 6.0 | 63.1 | 4.4 | 61.9 | 5.3 | 65.4 | 7.8 |
| Average number of years— | | | | | | | | |
| In current job [c] | 6.5 | 0.9 | 10.3*** | 0.9 | 10.8 | 1.1 | 9.3 | 1.8 |
| With positive earnings at ages 22–50 | 19.8 | 0.5 | 21.2* | 0.4 | 21.2 | 0.4 | 21.3 | 0.6 |
| Earnings and income ($) | | | | | | | | |
| Average— | | | | | | | | |
| Hourly wage | 14.0 | 0.8 | 16.4 | 1.0 | 15.1 | 1.1 | 19.3** | 1.9 |
| Annual earnings (self-reported) | 18,025 | 2,512 | 20,019 | 1,519 | 21,033 | 1,800 | 17,760 | 2,835 |
| Total household income | 45,652 | 3,879 | 56,266* | 3,453 | 55,134 | 4,094 | 58,789 | 6,472 |
| Annual earnings at ages 22–50 | 19,638 | 1,066 | 27,696*** | 1,083 | 26,317 | 1,199 | 30,577* | 2,258 |

SOURCE: Authors' calculations using HRS data linked to selected administrative data files from SSA.

NOTES: Sample consists of applicants aged 51 or older but younger than FRA who were denied at the initial level (including DDS reconsiderations).

Earnings and income values are adjusted to 2012 dollars using CPI.

All results are weighted to account for the complex HRS survey design and the varying probabilities of respondent consent to matching the HRS data to administrative data.

Chi-square tests compared the distribution of multinomial outcomes across groups; *t*-tests compared binomial outcomes.

* = statistically significant at the 0.10 level; ** = statistically significant at the 0.05 level; *** = statistically significant at the 0.01 level.

a. Indicators of statistical significance are relative to medical (step-2) denials.

b. Indicators of statistical significance are relative to step-4 denials.

c. Job held 2 years before DI application.

**Table 2.**
**Attributes of the preapplication job held by older DI applicants denied at the initial level, by reason for denial, 1992–2012 (in percent)**

| Characteristic | Denied for medical reasons (step 2) | | Denied for work capacity | | | | | |
|---|---|---|---|---|---|---|---|---|
| | | | Total (step 4 or 5) | | Able to perform past work (step 4) | | Able to perform other work (step 5) | |
| | Statistic | Standard error | Statistic [a] | Standard error | Statistic | Standard error | Statistic [b] | Standard error |
| Applicants with O*NET data (weighted) | | | | | | | | |
| Number | 581,770 | | 1,252,512 | | 890,475 | | 362,037 | |
| Percentage of total applicants | 13.1 | | 28.1 | | 20.0 | | 8.1 | |
| O*NET occupational requirement | | | | | | | | |
| General physical demands | | | | | | | | |
| Any | 38.5 | 5.8 | 43.1 | 4.0 | 36.1 | 4.5 | 58.2 | 7.6 |
| High-level | 3.8 | 2.3 | 2.9 | 1.3 | 3.2 | 1.7 | 2.2 | 2.2 |
| Flexibility or dexterity | | | | | | | | |
| Any | 32.4 | 5.6 | 28.6 | 3.6 | 23.6 | 4.0 | 39.3 | 7.5 |
| High-level | 4.3 | 2.4 | 12.8 | 2.7 | 9.2 | 2.7 | 20.6** | 6.2 |
| Vision | 10.4 | 3.6 | 15.2 | 2.9 | 17.3 | 3.6 | 10.6 | 4.8 |
| Cognitive ability demands | | | | | | | | |
| Any | 46.6 | 6.0 | 48.1 | 4.0 | 49.8 | 4.7 | 44.4 | 7.7 |
| High-level | 30.4 | 5.5 | 18.0 | 3.1 | 15.1 | 3.4 | 25.2 | 6.7 |
| Computer use | 15.4 | 4.3 | 23.4** | 3.4 | 25.7 | 4.1 | 18.5** | 6.0 |
| Interpersonal skills | 35.4 | 5.7 | 26.8 | 3.6 | 37.5 | 4.6 | 3.9*** | 3.0 |
| Stress tolerance | | | | | | | | |
| Any | 19.9 | 4.8 | 27.8** | 3.6 | 32.3 | 4.4 | 18.2 | 6.0 |
| High-level | 6.6 | 3.0 | 8.9 | 2.3 | 10.2 | 2.9 | 5.9 | 3.7 |
| O*NET occupational skill | | | | | | | | |
| Ability to— | | | | | | | | |
| Deal with unpleasant or angry people | 3.2 | 2.1 | 14.3*** | 2.8 | 10.8 | 2.9 | 21.8* | 6.4 |
| Use and update relevant knowledge | 8.3 | 3.3 | 8.9 | 2.3 | 4.3 | 1.9 | 18.6*** | 6.0 |
| Handle difficult working conditions | 13.2 | 4.0 | 23.0 | 3.4 | 17.0 | 3.6 | 35.9** | 7.4 |
| Skill type | | | | | | | | |
| Basic | 48.7 | 6.0 | 58.9* | 4.0 | 61.5 | 4.6 | 53.2 | 7.7 |
| Complex problem solving | 2.7 | 1.9 | 4.3 | 1.6 | 0.3 | 0.5 | 12.8*** | 5.2 |
| Resource management | 21.8 | 4.9 | 14.6 | 2.8 | 12.6 | 3.1 | 19.0 | 6.1 |
| Social | 43.9 | 5.9 | 20.6** | 3.3 | 22.8 | 4.0 | 15.8 | 5.6 |
| System | 17.1 | 4.5 | 5.7** | 1.9 | 2.4 | 1.5 | 12.8* | 5.2 |
| Technical | 13.2 | 4.1 | 17.3 | 3.0 | 11.0 | 3.0 | 30.8** | 7.1 |

SOURCE: Authors' calculations using HRS data linked to selected administrative data files from SSA.

NOTES: Sample consists of applicants aged 51 or older but younger than FRA who were denied at the initial level (including DDS reconsiderations).

Earnings and income values are adjusted to 2012 dollars using CPI.

All results are weighted to account for the complex HRS survey design and the varying probabilities of respondent consent to matching the HRS data to administrative data.

Chi-square tests compared the distribution of multinomial outcomes across groups; *t*-tests compared binomial outcomes.

* = statistically significant at the 0.10 level; ** = statistically significant at the 0.05 level; *** = statistically significant at the 0.01 level.

a. Indicators of statistical significance are relative to medical (step-2) denials.

b. Indicators of statistical significance are relative to step-4 denials.

postsecondary education (34 percent versus 57 percent). Further, step-4 denials were more likely than step-5 denials to work in large firms (26 percent versus 5 percent) but with a lower hourly wage rate ($15 versus $19) and lower average annual earnings at ages 22–50 ($26,317 versus $30,577).

Highlighting some intriguing differences in job attributes, Table 2 shows that step-4 denials were more likely than step-5 denials to work in an occupation requiring computer use (26 percent versus 19 percent) or interpersonal skills (38 percent versus 4 percent). However, they were less likely to have jobs that require high-level flexibility or dexterity (9 percent versus 21 percent) or the abilities to deal with unpleasant or angry people (11 percent versus 22 percent), to continually update their skills and knowledge (4 percent versus 19 percent), and to handle difficult working conditions (17 percent versus 36 percent). Interestingly, step-4 denials were also less likely to have jobs that demand complex problem-solving skills (0 percent versus 13 percent), system skills (2 percent versus 13 percent), or technical skills (11 percent versus 31 percent).

## How Postdenial Appeal, Reapplication, and Allowance Patterns Vary by Reason for Initial Denial

In the first of the two subsections that follow, we report our findings on the likelihood that DI applicants continue to pursue benefits after an initial denial. In the second subsection, we discuss how the availability of OASI benefits as early as age 62 may affect DI claiming behavior after initial denial.

### Applicants Who Appeal or Reapply for DI Benefits After Initial Denial

After the DDS denies an application at the initial or reconsideration level, applicants have several options for continuing to seek DI benefits (as well as the option not to continue). Chart 2 diagrams the options and shows how the older applicants who constitute our sample responded. The applicant must first decide whether to appeal the initial denial. For applicants who do not appeal, or whose appeal of the initial denial is likewise denied, the second decision is whether to reapply for benefits—perhaps several years after the initial denial. Overall, 56 percent of denied applicants in our sample ultimately received an award, of whom the majority had appealed the first observed denial. Some denied applicants (28 percent) did apply again, either after an appeal was denied or without first appealing.

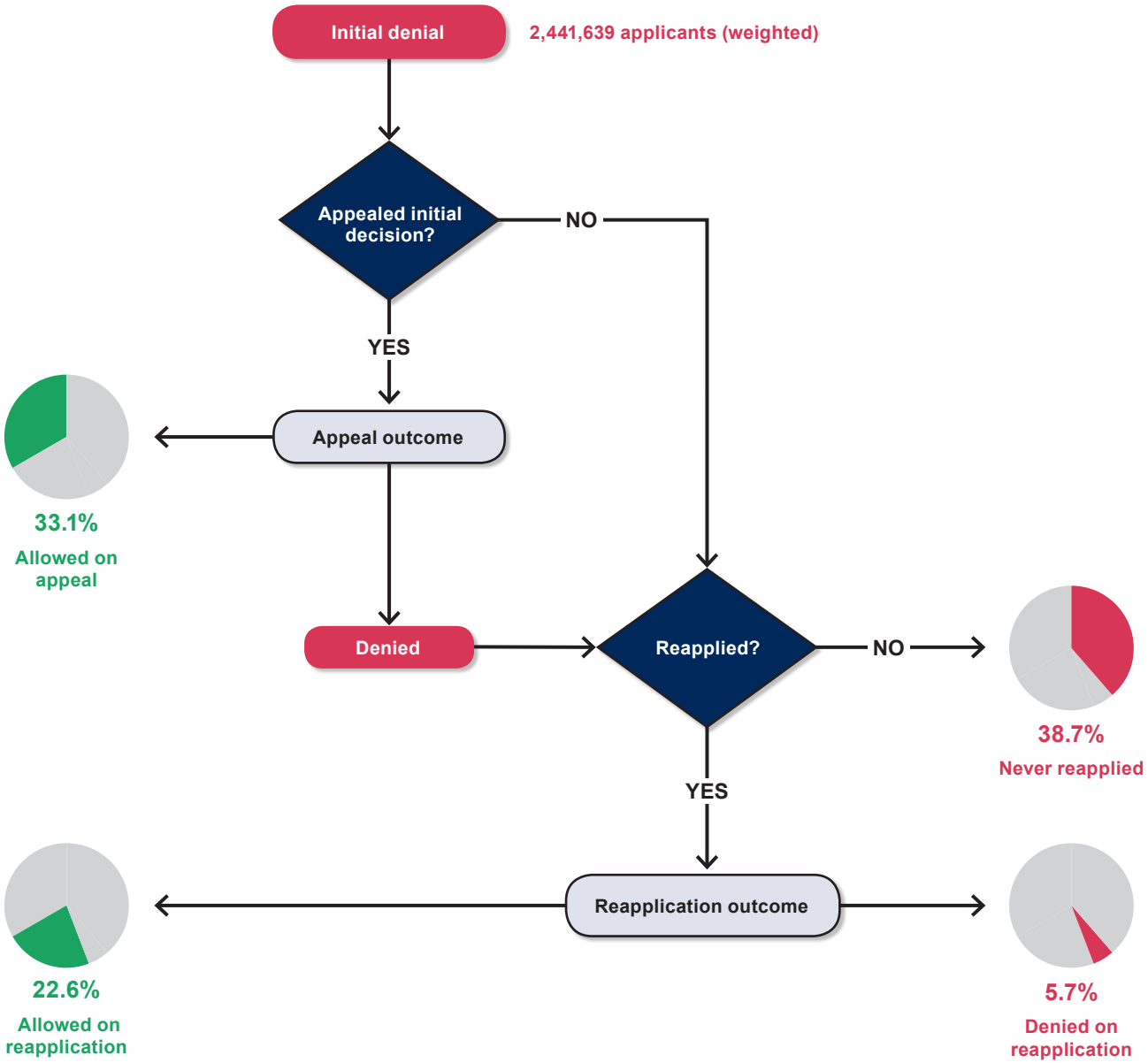
More than one-third of applicants with an initial-level denial (39 percent) never applied again, although that figure likely overstates the actual proportion for two reasons. First, the 831-file data prevent us from excluding two groups: individuals who appealed the initial decision but were denied; and applicants who appealed, were awarded DI benefits, then converted to OASI benefits on reaching FRA prior to May 2009. Second, as discussed earlier, some members of our sample may have reapplied after the period for which data were available. About 26 percent (weighted) had initial decisions within 3 years of the end of the period with administrative data available; within a longer follow-up period, we would expect more appeals or reapplications. Nonetheless, we have no reason to believe that the censoring would dramatically affect the differences in reapplication patterns across groups.

The overall pattern of appeals and reapplications among older applicants with initial-level denials reflects the important divergence in the paths of those denied for medical reasons and those denied for work-capacity reasons (Table 3). In particular, 67 percent of work-capacity denials were allowed after appeal or reapplication, more than twice the 31 percent rate at which those denied for medical reasons were ultimately allowed. Applicants with work-capacity denials have higher eventual DI allowance rates than applicants with medical denials have because a much higher share of the former successfully appeal the initial decision. Compared with the 20 percent of medical denials who appealed the initial decision and subsequently received benefits, 43 percent of those denied at step 4 and 32 percent of those denied at step 5 filed an appeal and received an allowance.

### Denied Applicants Who Claim OASI Before FRA

In addition to seeking DI benefits through appeal or reapplication, some denied applicants aged 62 or older may claim OASI benefits. For those who claim OASI before FRA, pursuing a DI application is relatively costless; SSA considers the DI eligibility for OASI claimants who report a health condition or impairment that might be significant enough to meet the DI criteria. Actuarially reduced OASI benefits claimed before FRA are an important component of the safety net for older adults with work-limiting health conditions (Leonesio, Vaughan, and Wixon 2003; Bound and Waidmann 2010; Schimmel and Stapleton 2012; Wu and Schimmel Hyde 2018). Even though OASI beneficiaries who are younger than FRA can earn more than the SGA

**Chart 2.**
**DI appeals and reapplications following initial denials for older applicants, 1992–2012**




SOURCES: Authors' calculations using HRS data linked to SSA's 831 file.

NOTES: Sample consists of applicants aged 51 or older but younger than FRA who were denied at the initial level (including DDS reconsiderations).

To meet the requirements for using restricted HRS-SSA linked data, we have rounded the percentages to avoid disclosing potentially identifiable information without diminishing the qualitative findings.

Percentages do not sum to 100.0 because of rounding.

**Table 3.**
**Postdenial appeals and reapplication for older DI applicants denied at the initial level, by reason for denial, 1992–2012 (in percent)**

| Characteristic | All denials | | Denied for medical reasons (step 2) | | Denied for work capacity | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | Total (step 4 or 5) | | Able to perform past work (step 4) | | Able to perform other work (step 5) | |
| | Statistic | Standard error | Statistic | Standard error | Statistic [a] | Standard error | Statistic | Standard error | Statistic [b] | Standard error |
| Number of applicants (weighted) | 2,441,638 | | 780,144 | | 1,661,494 | | 1,130,296 | | 531,198 | |
| Percentage of applicants who— | | | | | | | | | | |
| Appealed and were allowed | 33.1 | 2.4 | 19.8 | 3.6 | 39.4*** | 3.1 | 42.9 | 3.7 | 32.0*** | 5.6 |
| Reapplied and were— | | | | | | | | | | |
| Allowed | 22.6 | 2.2 | 11.2 | 2.9 | 28.0*** | 2.8 | 26.5 | 3.3 | 31.0*** | 5.6 |
| Denied | 5.7 | 1.2 | 8.1 | 2.5 | 4.5*** | 1.3 | 4.5 | 1.6 | 4.5*** | 2.5 |
| Never reapplied [c] | 38.7 | 2.5 | 60.9 | 4.4 | 28.3*** | 2.9 | 26.1 | 3.3 | 33.0*** | 5.7 |

SOURCE: Authors' calculations using HRS data linked to selected administrative data files from SSA.

NOTES: Sample consists of applicants aged 51 or older but younger than FRA who were denied at the initial level (including DDS reconsiderations).

All results are weighted to account for the complex HRS survey design and the varying probabilities of respondent consent to matching the HRS data to administrative data.

To meet the requirements for using restricted HRS-SSA linked data, we have rounded the percentages to avoid disclosing potentially identifiable information without diminishing the qualitative findings.

Chi-square tests compared the distribution of multinomial outcomes across groups.

*** = statistically significant at the 0.01 level.

a. Indicators of statistical significance are relative to medical (step-2) denials.

b. Indicators of statistical significance are relative to step-4 denials.

c. Includes applicants who appealed an initial denial and did not receive an allowance. The administrative data do not allow us to distinguish that group from initially denied applicants who neither appealed nor reapplied.

level and retain some benefits,[20] we hypothesize that denied DI applicants who ultimately claim retired-worker benefits before FRA are unlikely to reenter the workforce. Moreover, although pre-FRA OASI benefits offer a strong safety-net function and do not have the SGA restrictions that DI benefits do, applicants consider DI benefits to be financially preferable. Because OASI benefits claimed before FRA are actuarially reduced, a monthly OASI benefit claimed at age 62 will be up to 30 percent lower[21] than a monthly DI benefit at the same age, and the actuarial reduction will persist until death.

We find that a majority of applicants who were initially denied DI benefits go on to receive DI or OASI benefits before they reach FRA. Table 3 shows that approximately one-third of applicants denied for medical reasons and two-thirds of those denied for work-capacity reasons were observed to receive DI benefits by FRA. Among the cases for whom we did not observe a DI award, more than 70 percent began to receive either OASI or DI benefits at age 62 to FRA, with little difference across the groups of denied applicants (not shown).[22] Among those who began to receive benefits from age 62 to FRA, three-quarters received their first payment at age 62, suggesting that OASI claimed at the first possible age accounts for most of the benefits received among this group.

Most of the applicants in our sample were younger than the earliest OASI eligibility age when they applied for DI benefits, perhaps indicating that DI was their best option for income support following disability onset. Nonetheless, the ability to claim OASI at age 62 is relevant to individuals who apply for DI at ages nearing 62, as the initial decision takes months and appeals may take many months or years. Comparing step-4 and step-5 denials reveals that 68 percent of applicants denied at step 4 were aged from 57 to 61 when they applied, compared with 45 percent of applicants denied at step 5 (not shown). Our earlier findings indicated that many of these denied applicants appealed that decision, but for others, the ability to claim OASI benefits at age 62 may in part have driven a decision not to appeal.

A minority of applicants (12 percent) in our sample were aged 62 or older when they applied for DI benefits (not shown). Members of this group may have begun to receive OASI benefits while awaiting their DI decision, or they may have filed a claim for OASI benefits but SSA advised them that they might be eligible for DI benefits because of their limitations. There were slight, although not statistically significant, differences between denied and allowed DI applicants in the shares who were older than 62 at application.

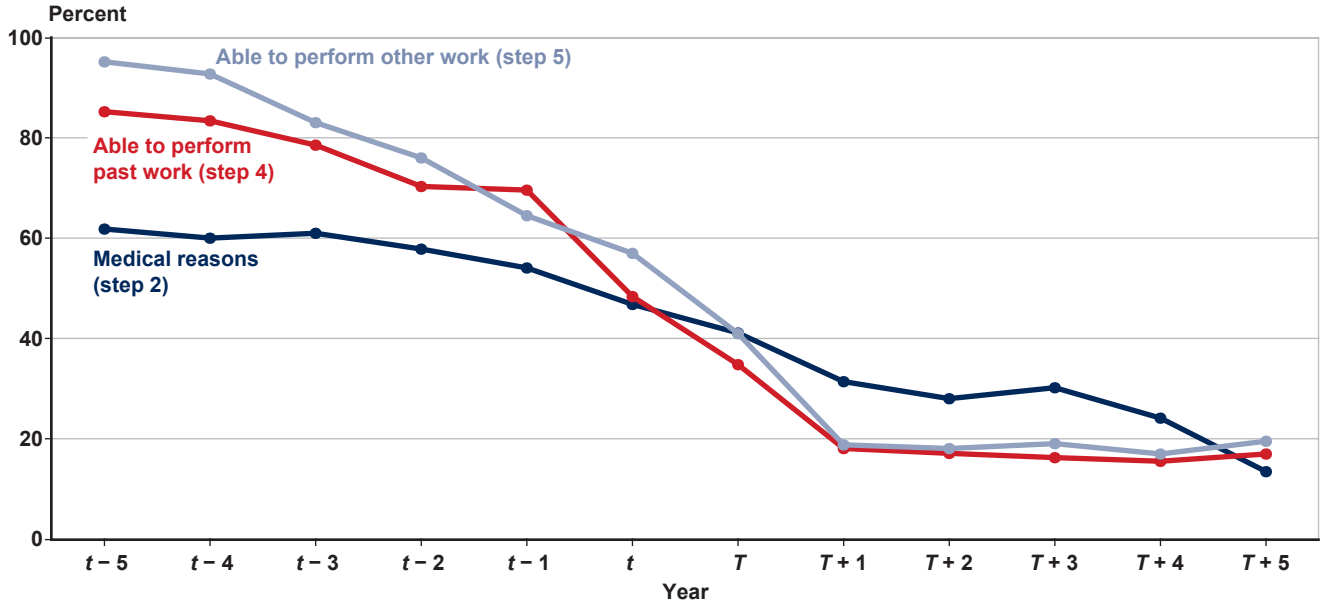
## *The Employment and Earnings Patterns of Denied Applicants Around the Time of the Application Decision*

We now turn to the employment and earnings trajectories of denied applicants in the 5 years before application and 5 years following the initial decision. In Charts 3 and 4, we indicate the application year with $t$ and the decision year with $T$; for many applicants, those events occur within the same year, and for them, earnings in $t$ and $T$ would thus be the same.

Five years before application, the employment rate among applicants who will be denied for work-capacity reasons is significantly higher than that of applicants who will be denied for medical reasons, although the employment rates of these groups converge by the year of application (Chart 3). In $t-5$, 85 percent of step-4 denials and 95 percent of step-5 denials were working, compared with 62 percent of those denied for insufficiently severe impairments at step 2 (that is, for medical reasons). As the application date approaches, the employment rates of all three groups drop, reaching 47 percent of step-2 denials, 48 percent of step-4 denials, and 57 percent of step-5 denials in the calendar year of application. The pre-application employment decline is steeper among those denied at steps 4 and 5 than for those denied at step 2. The preapplication employment rates we find for our sample resemble those in SSA (2017a) and Strand and Trenkamp (2015), despite differences in study-population age and timing before application. We omit confidence intervals to preserve the visual clarity of Chart 3; however, we note that the standard error on most of these estimates is relatively large (see Appendix Table A-1).[23] We are unable to reject the hypothesis that the likelihood of any earnings is the same in each year for applicants denied at steps 4 and 5. We are able to reject the hypothesis that the likelihood of any earnings is the same for medical and work-capacity denials, but only in the fifth and fourth years before denial; after that point, the likelihood of earnings is not statistically different across groups.

Although Chart 3 shows that the likelihood of working declines across all groups as the application date nears, the drop for those who will be denied for work-capacity reasons is more precipitous than that for applicants who will be denied at step 2. This pattern is most clearly seen when comparing employment rates 1 year before application ($t-1$) and 1 year after decision ($T+1$). Although the small (and declining) sample size for the postdecision years diminishes the precision of the estimates, applicants denied for

**Chart 3.**
**Employment rates (percentages of initially denied older applicants with positive earnings) in the calendar years before DI application and after denial, by reason for denial, 1992–2012**



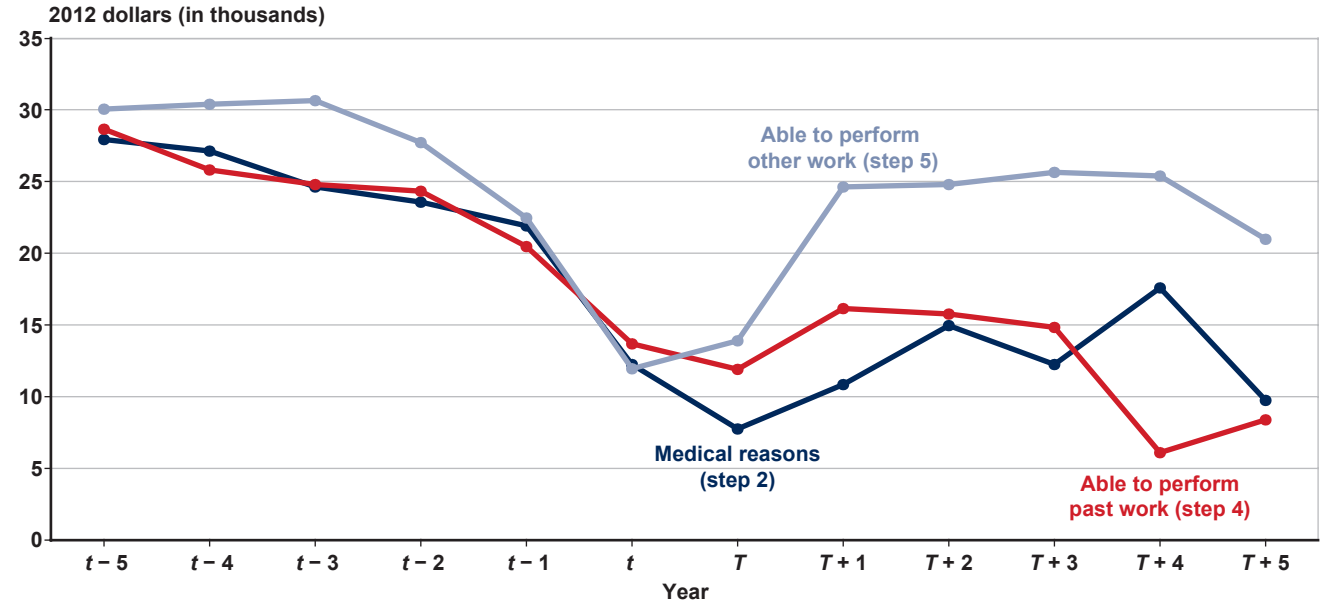

SOURCE: Authors' calculations using HRS data linked to selected administrative data files from SSA.

NOTES: *t* = year of application; *T* = year of initial denial. For many (but not all) applicants, *t* = *T*.

Sample consists of applicants aged 51 or older but younger than FRA who were denied at the initial level (including DDS reconsiderations).

See Appendix Table A-1 for tabulations, standard errors, and analogous results for various alternative combinations of covariates.

**Chart 4.**
**Mean earnings (among initially denied older applicants with positive earnings) in the calendar years before DI application and after denial, by reason for denial, 1992–2012**




SOURCE: Authors' calculations using HRS data linked to selected administrative data files from SSA.

NOTES: *t* = year of application; *T* = year of initial denial. For many (but not all) applicants, *t* = *T*.

Sample consists of applicants aged 51 or older but younger than FRA who were denied at the initial level (including DDS reconsiderations).

See Appendix Table A-1 for tabulations, standard errors, and analogous results for various alternative combinations of covariates.

medical reasons appear more likely to have earnings in the first few postdenial years than do those denied for work-capacity reasons. By the fifth calendar year after application, when the average age of a denied applicant in our study is from 61 to 63, 14–20 percent of sample members have positive earnings. This finding stands in contrast to Strand and Trenkamp's assessment: Observing younger denied applicants, they found that the likelihood of employment rose for the first 3 years after denial, before tapering off to a level that was similar to each group's employment rate a year before application.

Among older denied applicants who are working in a given year, we find that average earnings for those with step-2 and step-4 denials were much lower after denial than before application (Chart 4). Although applicants with step-5 denials had higher average earnings than other denied applicants, their postdenial earnings were about 15 percent lower than their preapplication averages. Given the small sample size and the small percentage of applicants who work after a denial, the standard errors for the earnings values are high and the results need to be interpreted with caution.[24] For example, applicants who were denied at step 5 and had observed earnings in $T+5$ numbered around a dozen.

The pool of individuals who return to work after DI denial may be subject to nonrandom self-selection; for example, those who continue to work after denial may have had above-average earnings prior to application and experienced a greater-than-average decline in earnings. The opposite could also be true. To assess this, we considered average (unadjusted) earnings in the years before application among those who worked in at least 1 of the 5 years after denial, and compared those values with average (unadjusted) preapplication earnings of the full group (not shown). No clear pattern emerged to suggest the direction of any selection into the group who returned to work. Applicants who worked after a step-4 denial generally had higher-than-average earnings in the years prior to application. Applicants who worked after denial at step 5 had lower-than-average earnings 3 to 5 years before application, but higher-than-average earnings in the application year and the 2 preceding years. Because we did not explore the hours worked among these applicants, we do not know whether those findings indicate individuals employed in higher-paying occupations or individuals working higher-than-average weekly hours before they applied.

The unadjusted values shown in Charts 3 and 4 do not account for demand-side or supply-side factors that might have affected the employment and earnings trajectories around the time of application. As shown in Tables 1 and 2, there were important differences that could explain the likelihood of returning to work after DI application, and we considered variants of our results that accounted for these factors. For example, applicants denied at step 4 were older on average than those denied at step 5, which could affect the likelihood of finding postdenial work. Results accounting for these observable differences across the groups are available in Appendix Table A-1. We present several variants of our results to show that they are relatively robust to specification changes. In particular, we highlight the changes resulting from the mechanism we use to account for group differences, and for more limited and fuller sets of controls.

In general, the pattern of regression-adjusted results is largely consistent with the pattern we describe based on the unadjusted values shown in Appendix Table A-1—employment rates decline for all denied-applicant subgroups and do not substantially recover in the 5 years after denial. Similar to the unadjusted mean employment rates, we cannot reject the hypothesis that the groups of denied applicants have the same employment rate in the regression-adjusted model in most instances. Average earnings continue to be volatile in a way that limits our ability to draw strong conclusions, and results vary across specifications. Further, if any individual characteristics are correlated with the reason for denial, holding characteristics constant across groups overcontrols for the factors that explain postdenial outcomes. Because the unadjusted results are more intuitive, they are our preferred specification.

## *Conclusions and Policy Implications*

We find that few older DI applicants who were initially denied benefits returned to work and that those who did return to work generally appeared to earn less on average than they had in the years before they applied. This general pattern is consistent with the findings of Strand and Trenkamp (2015), who focused on a younger population that had many more potential working years after a DI denial than did the denied applicants in our sample. As might be expected, we conclude that the likelihood of returning to work is lower among older denied applicants than among younger ones. We also find that older denied applicants exit the labor force earlier than their nonapplicant age-group peers.

Relatively few older workers go back to work following a denial, regardless of the reason for denial. For step-5 denials, disability examiners deemed the

applicants to be able to perform other work given their age, education, and work experience, and our descriptive analysis documents characteristics that are consistent with those criteria—they are younger, more educated, and working in what appear to be more skilled jobs than are applicants denied at step 4. We find evidence suggesting that applicants denied at step 5 who returned to work were slightly more likely to earn amounts that were similar to their preapplication earnings than were other applicant groups. Yet, small sample sizes lead us to interpret those results with caution.

After a denial, low employment among applicants appears to be related to appeal of the denial, DI reapplication, or filing an OASI claim. Among initially denied DI applicants aged 51 or older, virtually all who continued to seek benefits—whether DI or OASI—were receiving them by the time they reached FRA. Nearly two-thirds of the DI applicants in our sample who were initially denied for work-capacity reasons ultimately received a DI award after either appealing the initial decision or reapplying. That award rate is twice the rate for those who appealed or reapplied following a denial for medical reasons. Applicants who continue to pursue DI benefits have an incentive not to work at the SGA level, which would make them ineligible. Among denied applicants who do not receive DI benefits after a subsequent appeal or reapplication, more than two-thirds claim OASI benefits between the earliest age of eligibility (62) and their FRA. Although the work disincentives of OASI beneficiaries in this age range are less substantial than are those of DI beneficiaries, few of them work.

Thus, among the many older DI applicants who are initially denied benefits, few go back to work and most become Social Security disability or retirement program beneficiaries before attaining FRA. Two-thirds of DI applicants with initial work-capacity denials ultimately received DI benefits before FRA, and more than two-thirds of initially denied DI applicants who did not receive a subsequent DI allowance opted to receive actuarially reduced OASI benefits before reaching FRA. Many DI benefit allowances follow a lengthy appeal or reapplication. Because a DI award is predicated on not engaging in SGA, denied applicants awaiting a decision on an appeal are unlikely to work at significant levels and could experience financial hardship during a long wait. Over time, the size of SSA's backlog of appealed cases has varied, and one would expect an older denied applicant to be likelier to return to work when the backlog is shorter than when it is longer. This is because older workers who waited 2 to 3 years during a period of large backlogs for their appeal to be decided would have only a year or 2 after that decision and before FRA in which to reconsider work.

Two SSA demonstration projects test the feasibility of early interventions to support individuals with potential work-limiting medical conditions either after impairment onset (Retaining Employment and Talent After Injury/Illness Network, or RETAIN) or after a DI application is denied (Supported Employment Demonstration, or SED). Because the SED focuses on younger applicants (aged 18–50) with behavioral health issues, it is not directly relevant to the population we analyze. Yet the results of that demonstration may offer important insights about whether postdenial targeting of applicants is early enough to prevent labor force exit. RETAIN, which SSA is implementing jointly with the Department of Labor, will target workers after work-limitation onset but earlier in their trajectory toward benefit receipt. For older workers who may already be nearing labor force exit, this approach seems more likely to maintain potential connections to the labor force. However, the RETAIN demonstration is in its early stages and results are still many years away.

## Appendix A

Table A-1 presents regression model estimates of the conditional adjusted probabilities (in the case of employment) or means (in the case of earnings), based on the Stata statistical software package's *margins* command. Results are shown for each of four combinations of covariates, the first of which excludes all covariates (corresponding with the values plotted in Charts 3 and 4). For the second combination, the results are conceptually similar to those in a typical regression, where the group's own covariates are used to generate predicted probabilities. For the third and fourth combinations, we show the conditional adjusted value for each group assuming that it had the same average characteristics as the group of applicants denied at step 4. Stated another way, the conditional adjusted value for each group identifies the outcome if the average applicant who was denied at step 4 had been denied at step 2, 4, or 5, accounting for the covariates described in the narrative. The selection of group 4's characteristics was arbitrary, though the intent of holding constant at one group's characteristics was to minimize the effect of differences in characteristics across groups.

**Table A-1.**
**Employment and earnings of older DI applicants from 5 years before application to 5 years after initial decision, by initial outcome: Estimates using alternative combinations of covariates, 1992–2012**

| Year | No covariates [a] | | | | Using each group's average characteristics, full set of covariates | | | |
|---|---|---|---|---|---|---|---|---|
| | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error |
| | | | | ***Allowed for medical reasons*** | | | | |
| 5 years before application | 94.1 | 2.5 | 41,231 | 3,373 | 89.6 | 5.1 | 40,413 | 3,414 |
| 4 years before application | 93.7 | 2.6 | 41,038 | 3,574 | 89.6 | 5.1 | 38,960 | 3,303 |
| 3 years before application | 89.1 | 3.7 | 39,677 | 3,495 | 94.5 | 3.9 | 38,294 | 3,411 |
| 2 years before application | 84.1 | 4.2 | 39,949 | 2,972 | 88.0 | 5.4 | 35,942 | 3,558 |
| 1 year before application | 83.1 | 4.3 | 34,727 | 2,681 | 86.4 | 5.8 | 28,243 | 3,786 |
| Application year | 72.7 | 5.0 | 22,162 | 2,904 | 73.0 | 6.5 | 19,623 | 3,378 |
| Initial decision year | 57.3 | 5.6 | 19,042 | 3,351 | 48.4 | 8.1 | 15,770 | 2,990 |
| 1 year after initial decision | 17.6 | 4.2 | 13,405 | 4,330 | 16.1 | 5.9 | 15,246 | 6,480 |
| 2 years after initial decision | 13.5 | 3.6 | 15,553 | 4,863 | 14.0 | 5.8 | 24,183 | 4,442 |
| 3 years after initial decision | 14.4 | 3.9 | 14,355 | 5,446 | 17.1 | 6.9 | 15,462 | 7,651 |
| 4 years after initial decision | 12.7 | 3.7 | 4,848 | 1,442 | 5.7 | 4.5 | 13,456 | 3,034 |
| 5 years after initial decision | 8.4 | 3.2 | 5,928 | 2,055 | 8.6 | 5.7 | 12,047 | 2,515 |
| | | | | ***Allowed for medical-vocational reasons*** | | | | |
| 5 years before application | 96.9 | 1.5 | 38,524 | 1,988 | 95.0 | 2.7 | 32,631 | 2,229 |
| 4 years before application | 96.7 | 1.5 | 37,927 | 2,062 | 95.4 | 2.6 | 33,147 | 2,422 |
| 3 years before application | 93.7 | 2.0 | 35,477 | 2,144 | 95.2 | 2.3 | 29,663 | 2,678 |
| 2 years before application | 88.7 | 2.6 | 33,227 | 2,088 | 90.3 | 3.4 | 28,748 | 2,804 |
| 1 year before application | 83.3 | 3.0 | 28,890 | 2,146 | 85.4 | 3.8 | 24,997 | 3,560 |
| Application year | 67.4 | 3.6 | 17,145 | 2,181 | 58.1 | 5.2 | 22,446 | 5,887 |
| Initial decision year | 51.8 | 3.9 | 13,014 | 2,286 | 39.9 | 5.1 | 17,605 | 7,325 |
| 1 year after initial decision | 18.8 | 3.2 | 10,140 | 3,860 | 15.5 | 3.9 | 16,503 | 6,021 |
| 2 years after initial decision | 12.6 | 2.7 | 5,074 | 1,107 | 12.3 | 3.8 | 5,035 | 1,624 |
| 3 years after initial decision | 12.7 | 2.8 | 5,355 | 1,304 | 13.0 | 4.1 | 4,305 | 1,403 |
| 4 years after initial decision | 10.9 | 2.7 | 6,201 | 1,610 | 14.2 | 4.7 | 6,384 | 1,538 |
| 5 years after initial decision | 9.3 | 2.3 | 10,432 | 2,193 | 9.2 | 3.8 | 14,240 | 5,321 |
| | | | | ***Denied for medical reasons (step 2)*** | | | | |
| 5 years before application | 61.9 | 5.9 | 27,946 | 3,133 | 79.3 | 6.8 | 29,794 | 4,613 |
| 4 years before application | 60.1 | 5.9 | 27,137 | 3,003 | 74.8 | 6.8 | 30,714 | 4,859 |
| 3 years before application | 61.0 | 5.9 | 24,611 | 3,073 | 81.3 | 6.2 | 29,775 | 4,360 |
| 2 years before application | 57.9 | 5.9 | 23,551 | 3,094 | 80.5 | 6.3 | 26,285 | 4,300 |
| 1 year before application | 54.1 | 6.0 | 21,892 | 3,435 | 74.3 | 7.2 | 24,421 | 5,355 |
| Application year | 46.8 | 6.0 | 12,219 | 3,083 | 59.7 | 8.5 | 9,444 | 2,044 |
| Initial decision year | 41.2 | 6.0 | 7,737 | 957 | 53.9 | 8.5 | 7,904 | 1,241 |
| 1 year after initial decision | 31.4 | 5.9 | 10,828 | 2,593 | 40.8 | 9.3 | 13,826 | 3,551 |
| 2 years after initial decision | 28.0 | 6.0 | 14,934 | 3,134 | 36.4 | 8.8 | 18,777 | 2,766 |
| 3 years after initial decision | 30.2 | 6.5 | 12,250 | 3,623 | 33.2 | 9.3 | 18,409 | 3,584 |
| 4 years after initial decision | 24.2 | 6.0 | 17,601 | 4,526 | 26.8 | 8.9 | 18,891 | 3,882 |
| 5 years after initial decision | 13.5 | 4.4 | 9,731 | 2,146 | 10.2 | 5.8 | 9,407 | 1,697 |

(Continued)

**Table A-1.**
**Employment and earnings of older DI applicants from 5 years before application to 5 years after initial decision, by initial outcome: Estimates using alternative combinations of covariates, 1992–2012—*Continued***

| Year | No covariates [a] | | | | Using each group's average characteristics, full set of covariates | | | |
|---|---|---|---|---|---|---|---|---|
| | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error |
| | ***Denied for capacity to perform past work (step 4)*** | | | | | | | |
| 5 years before application | 85.3 | 3.6 | 28,642 | 2,454 | 82.5 | 6.7 | 24,286 | 4,696 |
| 4 years before application | 83.5 | 3.7 | 25,796 | 2,301 | 84.7 | 6.4 | 26,783 | 3,318 |
| 3 years before application | 78.6 | 4.1 | 24,806 | 2,505 | 85.6 | 5.9 | 25,734 | 4,031 |
| 2 years before application | 70.4 | 4.6 | 24,331 | 2,497 | 82.3 | 6.3 | 22,594 | 3,902 |
| 1 year before application | 69.6 | 4.5 | 20,469 | 2,656 | 75.6 | 6.2 | 17,465 | 5,609 |
| Application year | 48.4 | 5.0 | 13,694 | 3,004 | 36.8 | 7.4 | 10,353 | 2,901 |
| Initial decision year | 34.8 | 4.7 | 11,892 | 3,839 | 31.3 | 7.3 | 6,768 | 1,516 |
| 1 year after initial decision | 18.1 | 3.8 | 16,140 | 8,152 | 13.4 | 5.8 | 11,421 | 3,307 |
| 2 years after initial decision | 17.2 | 3.9 | 15,744 | 9,264 | 15.2 | 6.4 | 10,056 | 3,036 |
| 3 years after initial decision | 16.3 | 4.1 | 14,844 | 8,043 | 15.0 | 6.7 | 5,788 | 4,123 |
| 4 years after initial decision | 15.6 | 3.8 | 6,093 | 1,705 | 19.3 | 7.4 | 5,811 | 3,636 |
| 5 years after initial decision | 17.0 | 4.1 | 8,399 | 2,001 | 12.8 | 5.0 | 14,324 | 4,644 |
| | ***Denied for capacity to perform other work (step 5)*** | | | | | | | |
| 5 years before application | 95.2 | 2.5 | 30,061 | 3,721 | 96.9 | 2.3 | 22,353 | 3,938 |
| 4 years before application | 92.8 | 3.9 | 30,400 | 4,182 | 97.4 | 2.2 | 29,625 | 6,916 |
| 3 years before application | 83.1 | 6.2 | 30,668 | 5,063 | 91.5 | 8.3 | 33,036 | 7,268 |
| 2 years before application | 76.1 | 6.9 | 27,738 | 4,898 | 99.1 | 0.9 | 15,312 | 3,967 |
| 1 year before application | 64.5 | 7.7 | 22,470 | 4,526 | 88.5 | 7.5 | 10,861 | 3,629 |
| Application year | 57.0 | 7.8 | 11,931 | 2,999 | 65.1 | 11.8 | 7,420 | 4,019 |
| Initial decision year | 41.0 | 7.6 | 13,873 | 4,691 | 57.6 | 13.3 | 4,908 | 3,408 |
| 1 year after initial decision | 18.9 | 5.7 | 24,607 | 7,639 | 11.4 | 7.6 | 40,333 | 16,469 |
| 2 years after initial decision | 18.1 | 5.8 | 24,780 | 5,681 | 16.9 | 8.1 | 27,699 | 18,273 |
| 3 years after initial decision | 19.1 | 6.0 | 25,653 | 6,783 | 17.5 | 8.3 | 30,936 | 18,111 |
| 4 years after initial decision | 17.0 | 5.8 | 25,380 | 6,405 | 17.8 | 8.4 | 32,108 | 18,633 |
| 5 years after initial decision | 19.6 | 6.5 | 20,985 | 5,854 | 17.8 | 8.4 | 31,015 | 18,977 |

(Continued)

**Table A-1.**
**Employment and earnings of older DI applicants from 5 years before application to 5 years after initial decision, by initial outcome: Estimates using alternative combinations of covariates, 1992–2012—*Continued***

| Year | Using average characteristics of step 4 denials— | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | With full set of covariates | | | | Excluding covariates related to medical-vocational disability determinations | | | |
| | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error |
| ***Allowed for medical reasons*** | | | | | | | | |
| 5 years before application | 93.4 | 5.8 | 27,209 | 5,969 | 87.9 | 7.8 | 30,224 | 6,727 |
| 4 years before application | 93.4 | 5.8 | 26,231 | 5,775 | 87.2 | 8.0 | 30,003 | 6,732 |
| 3 years before application | 96.6 | 3.5 | 25,783 | 5,822 | 81.9 | 10.4 | 27,646 | 6,189 |
| 2 years before application | 92.3 | 6.7 | 24,199 | 5,662 | 73.6 | 12.5 | 27,408 | 6,258 |
| 1 year before application | 91.1 | 7.5 | 19,016 | 4,854 | 71.5 | 13.0 | 24,144 | 5,570 |
| Application year | 81.2 | 12.6 | 13,212 | 3,496 | 57.9 | 14.7 | 15,485 | 4,041 |
| Initial decision year | 59.7 | 18.9 | 10,618 | 2,937 | 39.6 | 14.1 | 14,463 | 4,007 |
| 1 year after initial decision | 23.7 | 14.9 | 10,265 | 4,813 | 11.4 | 6.3 | 8,696 | 3,354 |
| 2 years after initial decision | 20.8 | 14.1 | 16,282 | 4,686 | 8.1 | 4.8 | 11,490 | 4,141 |
| 3 years after initial decision | 25.0 | 16.0 | 10,410 | 5,614 | 8.7 | 5.1 | 10,077 | 4,498 |
| 4 years after initial decision | 9.1 | 9.2 | 9,060 | 2,407 | 7.3 | 4.5 | 3,032 | 1,256 |
| 5 years after initial decision | 13.4 | 11.9 | 8,111 | 1,945 | 5.5 | 3.6 | 3,708 | 1,702 |
| ***Allowed for medical-vocational reasons*** | | | | | | | | |
| 5 years before application | 97.3 | 2.5 | 23,729 | 4,413 | 93.5 | 4.6 | 26,392 | 5,786 |
| 4 years before application | 97.6 | 2.3 | 24,104 | 4,540 | 93.1 | 4.7 | 26,076 | 5,727 |
| 3 years before application | 97.4 | 2.3 | 21,571 | 3,967 | 87.4 | 7.2 | 24,885 | 5,551 |
| 2 years before application | 94.6 | 4.4 | 20,905 | 3,780 | 78.6 | 10.2 | 23,581 | 5,213 |
| 1 year before application | 91.5 | 6.5 | 18,178 | 3,602 | 72.6 | 11.7 | 19,968 | 4,556 |
| Application year | 70.3 | 15.9 | 16,323 | 4,232 | 52.3 | 14.1 | 12,655 | 3,263 |
| Initial decision year | 52.8 | 18.5 | 12,802 | 5,119 | 36.9 | 13.0 | 9,418 | 2,768 |
| 1 year after initial decision | 24.0 | 14.0 | 12,001 | 4,464 | 11.8 | 6.1 | 5,473 | 3,291 |
| 2 years after initial decision | 19.5 | 12.5 | 3,662 | 1,332 | 7.9 | 4.4 | 2,944 | 1,104 |
| 3 years after initial decision | 20.4 | 13.0 | 3,131 | 1,179 | 8.0 | 4.5 | 3,260 | 1,418 |
| 4 years after initial decision | 22.2 | 13.9 | 4,643 | 1,381 | 6.7 | 3.9 | 4,542 | 2,089 |
| 5 years after initial decision | 14.9 | 10.7 | 10,356 | 4,151 | 5.9 | 3.4 | 6,842 | 2,490 |
| ***Denied for medical reasons (step 2)*** | | | | | | | | |
| 5 years before application | 83.8 | 6.7 | 28,630 | 5,184 | 68.6 | 6.0 | 29,141 | 3,160 |
| 4 years before application | 79.7 | 7.0 | 29,513 | 5,533 | 66.6 | 5.9 | 29,099 | 3,165 |
| 3 years before application | 85.6 | 6.1 | 28,611 | 5,118 | 67.6 | 6.0 | 27,725 | 3,080 |
| 2 years before application | 84.9 | 6.3 | 25,257 | 4,948 | 64.0 | 6.1 | 25,758 | 3,230 |
| 1 year before application | 79.2 | 7.8 | 23,467 | 5,829 | 61.7 | 6.3 | 23,459 | 3,869 |
| Application year | 64.8 | 9.6 | 9,075 | 2,130 | 53.5 | 6.7 | 13,935 | 4,449 |
| Initial decision year | 58.8 | 9.5 | 7,595 | 1,278 | 46.6 | 6.9 | 7,467 | 1,032 |
| 1 year after initial decision | 45.0 | 10.6 | 13,285 | 3,476 | 37.5 | 7.1 | 11,176 | 2,653 |
| 2 years after initial decision | 40.3 | 9.8 | 18,044 | 3,084 | 34.1 | 6.8 | 15,535 | 2,884 |
| 3 years after initial decision | 36.9 | 10.2 | 17,690 | 3,786 | 32.1 | 7.1 | 14,425 | 3,922 |
| 4 years after initial decision | 30.0 | 9.4 | 18,153 | 4,237 | 28.9 | 6.5 | 17,351 | 4,066 |
| 5 years after initial decision | 11.6 | 6.3 | 9,039 | 1,911 | 17.0 | 5.7 | 8,385 | 1,728 |

(Continued)

**Table A-1.**
**Employment and earnings of older DI applicants from 5 years before application to 5 years after initial decision, by initial outcome: Estimates using alternative combinations of covariates, 1992–2012—*Continued***

| Year | Using average characteristics of step 4 denials— | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | With full set of covariates | | | | Excluding covariates related to medical-vocational disability determinations | | | |
| | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error | Percent with earnings [b] | Standard error | Mean of nonzero earnings ($) | Standard error |
| ***Denied for capacity to perform past work (step 4)*** | | | | | | | | |
| 5 years before application | 82.5 | 6.7 | 24,286 | 4,696 | 82.9 | 4.0 | 30,303 | 2,927 |
| 4 years before application | 84.7 | 6.4 | 26,783 | 3,318 | 80.9 | 4.1 | 27,845 | 2,901 |
| 3 years before application | 85.6 | 5.9 | 25,734 | 4,031 | 80.0 | 4.1 | 26,742 | 3,091 |
| 2 years before application | 82.3 | 6.3 | 22,594 | 3,902 | 74.0 | 4.5 | 24,213 | 2,844 |
| 1 year before application | 75.6 | 6.2 | 17,465 | 5,609 | 69.8 | 4.4 | 20,903 | 3,336 |
| Application year | 36.8 | 7.4 | 10,353 | 2,901 | 43.5 | 5.1 | 16,729 | 5,105 |
| Initial decision year | 31.3 | 7.3 | 6,768 | 1,516 | 33.1 | 4.9 | 14,671 | 6,291 |
| 1 year after initial decision | 13.4 | 5.8 | 11,421 | 3,307 | 20.4 | 4.3 | 18,060 | 10,118 |
| 2 years after initial decision | 15.2 | 6.4 | 10,056 | 3,036 | 16.4 | 4.3 | 25,852 | 12,222 |
| 3 years after initial decision | 15.0 | 6.7 | 5,788 | 4,123 | 15.8 | 4.5 | 20,750 | 9,116 |
| 4 years after initial decision | 19.3 | 7.4 | 5,811 | 3,636 | 15.4 | 4.1 | 5,333 | 1,375 |
| 5 years after initial decision | 12.8 | 5.0 | 14,324 | 4,644 | 16.9 | 4.5 | 7,490 | 1,800 |
| ***Denied for capacity to perform other work (step 5)*** | | | | | | | | |
| 5 years before application | 97.3 | 2.1 | 20,766 | 4,199 | 94.1 | 3.0 | 21,676 | 2,890 |
| 4 years before application | 97.7 | 1.9 | 27,523 | 7,045 | 91.1 | 4.3 | 20,900 | 4,444 |
| 3 years before application | 92.3 | 7.7 | 30,691 | 7,312 | 79.5 | 6.9 | 21,896 | 5,379 |
| 2 years before application | 99.2 | 0.8 | 14,225 | 3,996 | 71.4 | 7.8 | 18,399 | 4,183 |
| 1 year before application | 89.6 | 7.1 | 10,090 | 3,491 | 62.5 | 8.0 | 16,095 | 4,285 |
| Application year | 67.3 | 11.9 | 6,894 | 3,719 | 57.1 | 7.9 | 9,930 | 2,887 |
| Initial decision year | 60.0 | 13.3 | 4,559 | 3,103 | 44.6 | 8.0 | 12,548 | 4,677 |
| 1 year after initial decision | 12.4 | 7.8 | 37,470 | 14,247 | 21.6 | 6.1 | 25,834 | 7,539 |
| 2 years after initial decision | 18.4 | 8.2 | 25,733 | 16,096 | 20.4 | 6.0 | 21,870 | 4,912 |
| 3 years after initial decision | 19.0 | 8.5 | 28,741 | 15,840 | 20.9 | 6.2 | 19,849 | 6,291 |
| 4 years after initial decision | 19.4 | 8.5 | 29,829 | 16,252 | 18.6 | 6.0 | 17,511 | 6,123 |
| 5 years after initial decision | 19.4 | 8.5 | 28,813 | 16,616 | 22.4 | 7.0 | 15,567 | 4,613 |

SOURCE: Authors' calculations using HRS data linked to selected administrative data files from SSA.

NOTES: Values were developed in a generalized linear model regression with year, group, and year-by-group interactions and sampling weights applied.

Sample consists of applicants aged 51 or older but younger than FRA who were denied at the initial level (including DDS reconsiderations).

All results are weighted to account for the complex HRS survey design and the varying probabilities of respondent consent to matching the HRS data to administrative data.

SEF data are top-coded to the Social Security taxable maximum amount.

Earnings are adjusted to 2012 dollars using CPI.

The full set of covariates includes demand-side factors (the national employment rate in the specified year and the decision year); factors correlated with the decision to work (sex, race, Hispanic origin, marital status, self-reported health status, and self-reported functional limitations 2 years before application); and factors more closely related to the disability determination process, including age, education, and employment-related measures (firm size and job tenure in job held 2 years before application, and working-age years with positive earnings).

a. Values for denied applicants are plotted in Charts 3 and 4.

b. Limited to applicants with SEF data. Sample size and mean nonzero earnings decline in postdecision years because of right-censoring.

## Notes

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[1] Technically, the 30 percent maximum actuarial reduction will take effect in 2022, when the first of the workers whose FRA is 67 (those born in 1960 or later) turn 62. In 2020, the maximum reduction for claiming at age 62 is 28.3 percent.

[2] We use "initial denial" to refer to a DDS decision in either its first review or a first-level appeal known as a reconsideration. We exclude all subsequent appeals, which are heard by non-DDS entities (an administrative law judge, the Appeals Council, or a federal court).

[3] For simplicity, we use "work capacity" throughout the article to refer to step-4 and step-5 determinations, although that term is not SSA's official nomenclature.

[4] RFC is not considered if the individual has a severe impairment(s), has no past relevant work, is aged 55 or older, and has no more than a limited education. In such a case, the individual is simply deemed disabled (SSA 2018a).

[5] Since July 2012, SSA has allowed examiners to bypass step 4 for applicants whose past job requirements are difficult to determine from submitted information. For applicants found unable to adjust to another job in the economy in step 5, the examiners "return to the fourth step to develop the claimant's work history and make a finding about whether the claimant can perform his or her past relevant work" (SSA 2012a). This approach does not affect applicants during our period of study but may explain some of the differences in findings between our study and those of SSA's Office of the Inspector General, described later.

[6] More information about the OIS is available at https://www.ssa.gov/disabilityresearch /occupational_info_systems.html.

[7] The OIG report considered DI and Supplemental Security Income applicants aged 18–88 who were denied at steps 4 and 5 in 2013. Strand and Trenkamp focused on DI disabled-worker benefit applicants aged 18–61 who were initially denied at step 5 in 2005. Our study is limited to individuals aged 51 or older but younger than FRA who filed DI applications during 1992–2012.

[8] These linkages required permission from the HRS to access its restricted data, which was granted after receiving study approval from an independent Institutional Review Board under contract with Mathematica at the time of the study.

[9] For more information on O*NET, see Schimmel Hyde, Wu, and Gill (2018, Appendix B).

[10] Using the weights provided in the survey (HRS 2017), we estimate that our sample represents approximately 5.5 million individuals who ever applied for DI disabled-worker benefits after age 50 and before FRA from 1992 through 2012. Because the HRS consent weights were developed as cross-sectional weights at three points as new cohorts of HRS respondents entered the sample (1992, 1998, and 2004), we took the first weight available for each person, which effectively produces a sum of three cross sections over the period that are difficult to align with the flow of older applicants during our study period. Our estimated weighted sample size is significantly smaller than the 13.3 million DI applications filed between 1992 and 2012 by individuals older than 50 (calculated by SSA's Office of the Chief Actuary). SSA counts repeated applications from the same individual and we do not; this accounts for some of the difference in sample sizes. As we will show, repeated filings by older applicants may be common. Further, we restrict our sample to individuals who apply after the first HRS interview, which excludes applicants who filed before that interview (which can occur as late as age 56).

[11] The substance of our findings did not change between the weighted and unweighted versions of our analysis.

[12] Through the 2004 wave of the HRS, respondents who consented to the data linkage gave access to their records only through the year in which consent was granted. Since 2006, the HRS has asked respondents for prospective permission to link their records for 30 years. The HRS has obtained prospective permission from most respondents who previously offered permission as well as from the majority of new sample members, with the exception of those in early cohorts who were not reinterviewed after 2004.

[13] These respondents include 324 members of the 1931–1941 birth cohort (first interviewed by the HRS in 1992), 226 members of the 1942–1947 cohort (first interviewed in 1998), and 255 members of the 1948–1953 cohort (first interviewed in 2004). For respondents in the two oldest cohorts, we are able to observe all DI applications prior to the respondent's FRA (65 or 66, depending on birth year). For members of the 1948–1953 cohort, we are able to observe DI applications only through 2012, when those sample members were aged 59 to 64. For this latter cohort, we do not have complete information on OASI claiming, as none had reached FRA by the end of our observation period, and many had not reached age 62, the earliest age of OASI eligibility.

[14] We identified 242 respondents who applied for DI at least once before their first HRS interview but did not

subsequently reapply; we omitted those respondents from our analysis.

[15] Wixon and Strand excluded reconsideration determinations in their statistics. If their statistics had included reconsiderations, the difference between our allowance rates and theirs would be smaller, but the rate for our sample of older applicants would still be higher.

[16] More than 95 percent of our sample members had at least 1 year of earnings data in the SEF. We assumed that an individual did not have taxable earnings in any year with no SEF data. We estimate employment rates (the shares of individuals with earnings) as percentages of the sample members with at least 1 year of earnings data rather than percentages of the full sample.

[17] This issue is not exclusive to the CYBF; it applies to the Type of Claim field in the Master Beneficiary Record from which the CYBF is derived.

[18] System skills are defined by the Department of Labor's O*NET database as capacities used to understand, monitor, and improve sociotechnical systems. These include judgment and decision making, systems analysis, and systems evaluation.

[19] This finding is consistent with the medical-vocational grid's consideration of age, in that the likelihood of allowance, which can occur at step 5 but not at step 4, increases for older applicants.

[20] If an individual collects OASI benefits before FRA, his or her benefits are reduced by $1 for every $2 earned in excess of an annual limit (Song and Manchester 2007). In 2020, that limit is $18,240 annually. Any such benefit reductions result in actuarially fair increases in benefits paid to the beneficiary after attaining the FRA.

[21] See note 1.

[22] As we have discussed, the CYBF data do not allow us to identify whether those who receive an OASI or DI benefit after age 62 ultimately were awarded DI benefits or claimed OASI at an actuarially reduced amount. As an additional reason to interpret these statistics with caution, the data for 15 percent of our sample did not cover all years through age 62 and, for an additional 5 percent, the data extended beyond age 62 but not through FRA. With those cases excluded from consideration, virtually all of those whom we observe as able to claim OASI or DI benefits did so before FRA.

[23] For a version of this chart that includes the confidence intervals, see Schimmel Hyde, Wu, and Gill (2018, Figure V.1).

[24] For a version of this chart that includes the confidence intervals, see Schimmel Hyde, Wu, and Gill (2018, Figure V.2).

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