However, the econometric application of the piecewise linear budget constraint method has been called into question by ... is defined—only if the estimated coefficients yield a positive compensated substitution effect. When this condition was not satisfied, researchers imposed it by constraining the income coefficient to be negative. MaCurdy ... is not warranted. The compensated effect may be estimated to be positive without the researcher imposing it.

Note: In 1983 the earnings test was eliminated for 70–71 year olds (71–72 year olds in the following March CPS) but was not changed for 62–69 year olds. See Figure 2 note.

However, the econometric application of the piecewise linear budget constraint method has been called into question by researchers. Only if the estimated coefficients yield a positive compensated substitution effect does the constraint hold—when this condition was not satisfied, researchers imposed it by constraining the income coefficient to be negative. MaCurdy et al. (1992) note that imposing the constraint is not warranted. The compensated effect may be estimated to be positive without the researcher imposing it, and this is observed in Figure 3-A through 3-D.

Note: In 1983 the earnings test was eliminated for 70–71 year olds (71–72 year olds in the following March CPS) but was not changed for 62–69 year olds. See Figure 2 note.

Source: Friedberg (2000), p. 56
Consumption-Smoothing Benefits of Social Security

Living Standards of the Elderly

**FIGURE 13-2**

**Elderly Poverty and Social Security, 1959–2004** • There is a striking negative correspondence over time between the poverty rates of the elderly (which have fallen) and the size of the Social Security program (which has risen).

Figure 4. Change in consumption at retirement, by wealth quartile

Source: Bernheim et al. (2001), p. 847
Fig 1.—Percentage change in food expenditure, predicted food consumption index, and time spent on food production for male household heads by three-year age ranges. Data are taken from the pooled 1989–91 and 1994–96 cross sections of the CSFII, excluding the oversample of low-income households. The sample is restricted to male household heads (1,510 households). All series were normalized by the average levels for household heads aged 57–59. All subsequent years are the percentage deviations from the age 57–59 levels. See Sec. IV for details of data and derivation of food consumption index.

Source: Aguiar and Hurst (2005), p. 925
Evidence

**FIGURE 13-3**

**Elderly Work and Social Security, 1959–2004** • There is a striking negative correspondence over time between the labor force participation (LFP) rates of the elderly (which have fallen) and the size of the Social Security program (which has risen).

**Evidence**

**Figure 13-4**

**Hazard Rate of Retirement for Males in the United States**

- The male hazard rate, or exit rate at each age given that a man has worked to that age, has a distinct spike at age 62 (the Early Entitlement Age, EEA) and 65 (the Full Benefit Age, FBA), key ages for the Social Security system.

*Source: Diamond and Gruber (1999), Figure 11.12.*

---

**retirement hazard rate** The percentage of workers retiring at a certain age.
13.3 Social Security and Retirement

Evidence

**FIGURE 13-5**

The Evolution of the U.S. Male Retirement Hazard

- In 1960, before the EEA of 62 was introduced for men, the hazard rate for men was highest at age 65 (the FRA), with no spike at age 62. By 1970, the spike at 62 had begun to emerge, and by 1980 it was larger than the spike at age 65.

Source: Gruber and Wise (1999), Figure 12
Social Security and Retirement

Evidence

**FIGURE 13-6**

Hazard Rate of Retirement in France

In France, there is an enormous exit rate from the labor force at age 60, which is both the EEA and FBA.

Source: Gruber and Wise (1998, Figure 11).
Social Security and Retirement

Evidence

**FIGURE 13-7**

Change in Average Retirement Age in Germany from 1968 to 1992

- Germany lowered its age of social insurance entitlement by five years (from 65 to 60) in 1973; within seven years, the average age at which individuals retire had fallen from 63 to 58.

Source: Gruber and Wise (1999), Figure 5.
Implicit Social Security Taxes and Retirement Behavior

**APPLICATION**

**FIGURE 13-8**

**Implicit Taxes on Work and Nonwork**

There is large variation across nations in the social security disincentives to work at older ages. The disincentive to work is measured here as the natural logarithm of the sum of implicit taxes on work at older ages. Those nations with greater disincentives to work tend to have much higher nonwork among older workers.

Source: Gruber and Wise (1999), Figure 17.
The potential for omitted-variable bias, though, is more important. For example, with out adequate controls for macro shocks, the effect of Social Security on GDP may be imperfect, and so some positive bias presum-
ably remains. A negative \( \beta \) hypothesis that the effect of Social Security on GDP is actually positive may be offset by the early 1980s and perhaps early 1990s, when Social Security surpluses might appear positively correlated with GDP growth.

Statistically significant at the 2-percent level.

Table 1—Least-Squares Regression with (Robust) Standard Errors, 1949–2002 (Dependent Variable: Modified Primary On-Budget Surplus, \( S^{\text{ON}}_t \))

<table>
<thead>
<tr>
<th>Variable</th>
<th>Specification</th>
</tr>
</thead>
<tbody>
<tr>
<td>( S^{\text{OFF}}_t )</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>(0.736)</td>
</tr>
<tr>
<td>GDP(_t)</td>
<td>2</td>
</tr>
<tr>
<td></td>
<td>(0.688)</td>
</tr>
<tr>
<td>Year (( t ))</td>
<td>3</td>
</tr>
<tr>
<td></td>
<td>(0.094)</td>
</tr>
<tr>
<td>Year(^2) (( t^2 ))</td>
<td>4</td>
</tr>
<tr>
<td></td>
<td>(0.000012)</td>
</tr>
<tr>
<td>Wages and salaries</td>
<td></td>
</tr>
<tr>
<td>Intercept term</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.002)</td>
</tr>
<tr>
<td>( R^2 )</td>
<td></td>
</tr>
</tbody>
</table>

Notes: Robust standard errors shown in parentheses.

\( \dagger \) Statistically significant at the 2-percent level.

Source: Smetters (2004), p. 179
Table 2—Least-Squares Regression with (Robust) Standard Errors, Different Periods (Dependent Variable: Modified Primary On-Budget Surplus, $S_{t}^{ON}$)

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>$S_{t}^{OFF}$</td>
<td>-0.476</td>
<td>-2.83\textsuperscript{†}</td>
</tr>
<tr>
<td></td>
<td>(1.413)</td>
<td>(0.73)</td>
</tr>
<tr>
<td>GDP\textsubscript{t}</td>
<td>0.208</td>
<td>0.059</td>
</tr>
<tr>
<td></td>
<td>(0.157)</td>
<td>(0.162)</td>
</tr>
<tr>
<td>Year ((t))</td>
<td>0.001</td>
<td>-0.003</td>
</tr>
<tr>
<td></td>
<td>(0.003)</td>
<td>(0.004)</td>
</tr>
<tr>
<td>Year$^2$ ((t^2))</td>
<td>-0.00003</td>
<td>0.00005</td>
</tr>
<tr>
<td></td>
<td>(0.00009)</td>
<td>(0.00004)</td>
</tr>
<tr>
<td>Wages and salaries</td>
<td>-0.161</td>
<td>0.737\textsuperscript{†}</td>
</tr>
<tr>
<td></td>
<td>(0.233)</td>
<td>(0.25)</td>
</tr>
<tr>
<td>Intercept term</td>
<td>-0.133</td>
<td>-0.417\textsuperscript{†}</td>
</tr>
<tr>
<td></td>
<td>(0.114)</td>
<td>(0.14)</td>
</tr>
</tbody>
</table>

Notes: Standard errors shown in parentheses. Test for structural break at 1970 (Wald test): $P = 0.004$.

\textsuperscript{†} Statistically significant at the 2-percent level.
Beneficiaries in Current-Payment Status

Chart 2. All Social Security disabled beneficiaries in current-payment status, December 1970–2017

The number of disabled beneficiaries has risen from 1,812,786 in 1970 to 10,059,166 in 2017, driven predominately by an increase in the number of disabled workers. The number of disabled adult children has grown slightly, and the number of disabled widow(er)s has remained fairly level. In December 2017, there were 8,695,475 disabled workers; 1,105,405 disabled adult children; and 258,286 disabled widow(er)s receiving disability benefits.

SOURCE: Table 3.
The percentage of disabled-worker beneficiaries increases with age for both men and women. In December 2010, the largest percentage of disabled-worker beneficiaries was aged 60–64. Disability benefits convert to retirement benefits when the worker reaches full retirement age, 65–67, depending on the year of birth.

Source: SSA DI annual report

NOTE: FRA = full retirement age.
Chart 8.  
Social Security disability awards, 1980–2010

The total number of awards decreased from 1980 through 1982, started to rise in 1983, and began to increase more rapidly in 1990. Awards for disabled-worker benefits have been most pronounced and drive the overall pattern shown in the total line. They increased from a low of 297,131 in 1982 to 636,637 in 1992, were relatively flat from 1992 through 2000, and started to increase again in 2001. There were 1,026,988 worker awards in 2010. Other awards have risen at a much slower rate. Awards to disabled adult children have gradually increased from 33,470 in 1980 to 81,681 in 2010. Awards to disabled widow(er)s have risen from just over 16,000 in 1980 to 33,259 in 2010.

SOURCE: Table 35.
In 2010, 1,026,988 disabled workers were awarded benefits. Among those awardees, the most common impairment was diseases of the musculoskeletal system and connective tissue (32.5 percent), followed by mental disorders (21.4 percent), circulatory problems (10.2 percent), neoplasms (9.0 percent), and diseases of the nervous system and sense organs (8.2 percent). The remaining 18.7 percent of awardees had other impairments.

SOURCE: Table 37.

a. Data for individual mental disorder diagnostic groups are shown separately in the pie chart below.
Nonparticipation and Recipiency Rates, Men 45-54 Years Old

Source: Parsons 1984 Table A1
Table 1—Reassessments of Initial Social Security Determinations

A. Bureau of Disability Insurance Review One Year After Initial Determination (Percentages):

<table>
<thead>
<tr>
<th>BDI assessment</th>
<th>Allowance</th>
<th>Denial</th>
</tr>
</thead>
<tbody>
<tr>
<td>Allowance</td>
<td>78.8</td>
<td>21.1</td>
</tr>
<tr>
<td>Denial</td>
<td>22.5</td>
<td>77.5</td>
</tr>
</tbody>
</table>

Note: The sample sizes are 250 initial allowances and 248 initial denials.

<table>
<thead>
<tr>
<th></th>
<th>1972</th>
<th>1978</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Population</td>
<td>Rejected Applicants</td>
</tr>
<tr>
<td>Labor Supply</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Percent Employed</td>
<td>77.7</td>
<td>32.6</td>
</tr>
<tr>
<td>Percent Worked 71/77</td>
<td>91.9</td>
<td>45.0</td>
</tr>
<tr>
<td>Percent Full Year (≥ 50 Weeks)²</td>
<td>76.8</td>
<td>47.4</td>
</tr>
<tr>
<td>Percent Full Time (≥ 35 Hours)²</td>
<td>95.4</td>
<td>75.9</td>
</tr>
<tr>
<td>Earnings Among Positive Earners</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Median Annual Earnings, 71/77</td>
<td>$9000</td>
<td>$4000</td>
</tr>
</tbody>
</table>

Source: Bound 1991
length of the unit interval: 0.02 to 1. However, only a few examiners have such extreme allowance rates: the first and ninety-ninth percentiles of $EXALLOW$ are 0.17 and 0.64, respectively.\(^3\)\(^8\) Figure 3 presents smoothed histograms at the examiner level of examiners’ deviations from the mean initial allowance rate in their DDS office, unadjusted and regression-adjusted for differences in case mix. Case controls include the fraction of cases in each of nine age bands, 14 body system codes, alleged terminal illness, three-digit zip code, and decision month, as well as a variable measuring average prior earnings of the set of applicants assigned to a given examiner. Adjusting for case mix reduces variation in initial allowance rates, but there is still significant variation remaining (the standard deviation is 0.06, compared with 0.10 unadjusted).

Two key assumptions underlie our empirical strategy. First, in order for $EXALLOW$ to be a valid instrument for SSDI receipt, applicants’ assignment to DDS examiners must be uncorrelated with unobserved characteristics such as impairment severity conditional on observed characteristics. This amounts to an assumption of conditional random assignment to DDS examiner within a DDS. That is, at most, examiners may specialize in a particular type of impairment (e.g., mental disorders) or age group, but within this type, examiners do not further specialize in cases of either low or high severity. As discussed previously, applicants are assigned to

\(^{38}\) Despite the fact that we condition on examiners with caseloads of 30 or more, one might be concerned that examiners with relatively few observations will tend to have very high or very low allowance rates because they are noisier. We explored this possibility by applying a Bayesian “shrinkage” estimator to $EXALLOW$ (see, e.g., Kane and Staiger 2008) and estimating our results using this “corrected” instrument. The new instrument had a range of 0.14 to 0.75. Both the first and second stage (labor supply) estimates were slightly higher using this alternative instrument, but not significantly so, and the patterns in the coefficients remained the same.
for stratification of examiners across DDS offices. We display $t$-statistics in parentheses, where robust standard errors are computed and clustered by DDS examiner. Column 1 shows the first-stage coefficient on $EXALLOW$ from a regression with no additional covariates. In both years, a 10 percentage point increase in initial examiner allowance rate leads to an approximately 3 percentage point increase in the probability of ultimately receiving SSDI.

Adding covariates sequentially to the regression allows us to indirectly test for random assignment on the basis of observable characteristics because only covariates that are correlated with $EXALLOW$ will affect the estimated coefficient on $EXALLOW$ when included. Based on our interviews with DDS managers (see Section I), we expect the additions of the body system and terminal illness indicators to potentially affect the coefficient on $EXALLOW$, since they are case assignment variables, but no other variables should affect the coefficient. The coefficient on $EXALLOW$ falls from 0.29 to 0.24 with the addition of body system codes and is not significantly affected by the addition of any other variables, including the TERI flag. Thus, our results are consistent with random assignment of applicants to examiners within DDS office, conditional on body system code and alleged terminal illness.$^{40}$

$^{40}$We also experimented with a different measure of initial allowance rate to test the implication of the monotonicity assumption that generic allowance rates can be used to instrument for any type of case. For this measure, we constructed the initial allowance rate leaving out all cases with the same body system code as the applicant (instead of just the applicant’s own case). Table A1 in the online Appendix presents these results. For all impairments but one (“special/other” cases, around 4 percent of the sample), this alternative measure of $EXALLOW$ is positively and significantly associated with increased SSDI receipt. (We replicated our analysis of labor supply effects dropping this
the predicted probability of SSDI receipt. Specifically, we regress initial allowance decisions on indicators for type of impairment, age group, decision month, and DDS, as well as a measure of average prior earnings, and construct the residual, $Z$, which by construction is orthogonal to the case mix controls and varies systematically only with $EXALLOW$. Then we estimate a probit of ultimate SSDI receipt on the residualized $Z$. This is our measure of the predicted probability of SSDI receipt, $P(Z)$. Next we estimate a local quadratic regression of employment on predicted SSDI receipt and compute the numerical derivative of this function to estimate $\frac{\partial E[y]}{\partial P(Z)}$.

Figure 7 shows the MTE as a function of unobserved severity, where severity is reverse ordered and measured in percentiles (see definition of $u$ in Section IVA), along with boot-strapped 95 percent confidence intervals. Applicants on the margin for an examiner with a predicted SSDI receipt rate of 65 percent (the mean rate) are in the sixty-fifth percentile of the unobserved (reverse) severity distribution. That is, they have an impairment that is less severe than 65 percent of applicants, and more severe than 35 percent of applicants. Since we estimate that 57 percent of applicants are always takers (that is, they would receive SSDI benefits regardless of initial examiner assignment), the MTE is not identified for applicants on the margin of SSDI receipt rates less than 57 percent. Similarly, the MTE is not identified for applicants on the margin of SSDI receipt rates greater than 80 percent ($= 57 + 23$, the fraction of marginal applicants). As a result, we are only able to trace the MTE for applicants between the fifty-seventh and eightieth percentiles of the unobserved (reverse) severity distribution (or the twenty to forty-third percentiles of the actual unobserved severity distribution $s$). The estimates become imprecise at the more extreme ends of the distribution since there are relatively small numbers of examiners with margins at these points.

**Figure 7. Marginal Treatment Effect on Employment**

*Notes:* Ninety-five percent confidence intervals shown with dashed lines. Bandwidth is 0.084.

Coefficient = -0.849, se = 0.164 t = -5.18

Source: Autor and Duggan 2003
Fig. 1. Early Retirement Ages by Pension Type

Notes: The vertical lines mark the beginning of changes implemented under the 2000 and 2004 pension reforms.

Source: Manoli and Weber '13
Fig. 2. Pre-Reform Pension Claims & Job Exits

Notes: For computing the survival curves, the sample is restricted to pre-reform birth cohorts (1930 through 1939 for men and 1935 through 1944 for women) and also to individuals for whom a claim is observed prior to age 70. See Table 1 for the full sample restrictions.

Source: Manoli and Weber '13
Fig. 5A. Men’s Claiming Ages & Exit Ages by Cohort

Source: Manoli and Weber '13
Fig. 5A. Men’s Claiming Ages & Exit Ages by Cohort

Source: Manoli and Weber '13
Fig. 5A. Men’s Claiming Ages & Exit Ages by Cohort

Source: Manoli and Weber ‘13
Fig. 5A. Men’s Claiming Ages & Exit Ages by Cohort

Source: Manoli and Weber '13
Fig. 9. Claiming & Exiting by Birth Cohort & Contribution Years, Men

Notes: Each figure plots the fraction individuals still in the labor market who claim pensions or exit jobs by birth cohort. Women with 40 or more contribution years and men with 45 or more contribution years are exempt from the increases in the Early Retirement Ages and can continue to retire at ages 55 and 60 respectively. The sample is restricted to men ages 59 through 62 in birth cohorts 1939 through 1947 and women ages 54 through 57.75 in birth cohorts 1944 through 1952. Observations are censored at the Early Retirement Age specified for each individual.
Fig. 9. Claiming & Exiting by Birth Cohort & Contribution Years, Women

Notes: Each figure plots the fraction individuals still in the labor market who claim pensions or exit jobs by birth cohort. Women with 40 or more contribution years and men with 45 or more contribution years are exempt from the increases in the Early Retirement Ages and can continue to retire at ages 55 and 60 respectively. The sample is restricted to men ages 59 through 62 in birth cohorts 1939 through 1947 and women ages 54 through 57.75 in birth cohorts 1944 through 1952. Observations are censored at the Early Retirement Age specified for each individual.

Source: Manoli and Weber '13
Notes: There are two forms of government-mandated retirement benefits in Austria: (1) government-provided pension benefits and (2) employer-provided severance payments. The employer-provided severance payments are made to private sector employees who have accumulated sufficient years of tenure by the time of their retirement. Tenure is defined as uninterrupted employment time with a given employer and retirement is based on claiming a government-provided pension. The payments must be made within 4 weeks of claiming a pension according to the following schedule. If an employee has accumulated at least 10 years of tenure with her employer by the time of retirement, the employer must pay one third of the worker's last year's salary. This fraction increases from one third to one half, three quarters and one at 15, 20 and 25 years of tenure respectively. Since payments are based on an employee's salary, overtime compensation and other non-salary payments are not included when determining the amounts of the payments. Provisions to make these payments come from funds that employers are mandated to hold based on the total number of employees. Severance payments are also made to individuals who are involuntarily separated (i.e. laid off) from their firms if the individuals have accumulated sufficient years of tenure prior to the separation. The only voluntary separation that leads to a severance payment, however, is retirement. Employment protection rules hinder firms from strategically laying off workers to avoid severance payments and there is no evidence on an increased frequency of layoffs before the severance pay thresholds.

Source: Manoli and Weber NBER'11
Fig. 3. Distribution of Tenure at Retirement, Full Sample

Notes: This figure plots the distribution of tenure at retirement at a monthly frequency. Each point captures the number of people that retire with tenure greater than the lower number of months, but less than the higher number of months. Tenure at retirement is computed using observed job starting and job ending dates. Since firm-level tenure is only recorded beginning in January 1972, we restrict the sample to individuals with uncensored tenure at retirement (i.e. job starting after January 1972).

Source: Manoli and Weber NBER'11
Fig. 6. Tenure at Retirement by Health Status

Notes: Health status is measured based on the fraction of time between age 54 and retirement that is spent on sick leave. An individual is classified as unhealthy if his health status is below the median level. The median health status is computed within the sample of individuals with positive sick leave and uncensored tenure at retirement.; this median health status is 0.076.

Source: Manoli and Weber NBER'11
Figure E.6: Adjustment Across Ages: Histograms of Earnings and Normalized Excess Mass, 59-73-year-olds Claiming OASI by Age 65, 2000-2006

Panel A: Earnings histograms, by age

Panel B: Normalized excess mass, by age

See notes to Figure 2. The figure differs from Figure 2 only because the years examined are 2000-2006 (whereas in Figure 2 the years examined are 1990-1999). As explained in the main text, the NRA slowly rose from 65 for cohorts that reached age 62 during this period; the results are extremely similar when the sample is restricted to those who claimed by 66, instead of 65. In the year of attaining NRA, the AET applies for months prior to such attainment.
Mandated Savings (M) Around Eligibility Threshold in 1998

Source: Chetty et al. QJE 2014
Effect on Mandate on Total Pension Contributions

Source: Chetty et al. QJE 2014
Effect on Mandate on Total Pension Contributions

Percent with Total Pension Contribution > DKr 1265

Income (DKR 1000s)

Empirical

Total Pensions Pass-Through Rate: $\phi_G = 85\%$ (11%)

Source: Chetty et al. QJE 2014
Effect on Mandate on Total Saving

![Graph showing the effect on mandate on total saving. The graph plots the percent with total savings greater than DKr 1371 against income (DKR 1000s). The graph includes an empirical line and a predicted line with 100% pass-through rate.]

Total Pensions Pass-Through Rate: \( \phi_G = 127\% \)

Source: Chetty et al. QJE 2014
### Mandated Savings Plan: Pass-Through Estimates

<table>
<thead>
<tr>
<th>Dep. Var.:</th>
<th>Δ Total Pensions</th>
<th>Total Pension Threshold</th>
<th>Total Saving Threshold</th>
<th>Total Ind. Saving Threshold</th>
<th>Net Saving Threshold</th>
</tr>
</thead>
<tbody>
<tr>
<td>Pass-Through Estimate</td>
<td>0.883 (0.204)</td>
<td>0.845 (0.113)</td>
<td>1.268 (0.363)</td>
<td>1.336 (0.349)</td>
<td>2.188 (0.587)</td>
</tr>
<tr>
<td>Research Design</td>
<td>Linear</td>
<td>Linear</td>
<td>Quadratic</td>
<td>Linear</td>
<td>Linear</td>
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<tr>
<td>No. of Obs</td>
<td>35,578</td>
<td>35,578</td>
<td>35,578</td>
<td>158,229</td>
<td>148,380</td>
</tr>
</tbody>
</table>

Source: Chetty et al. QJE 2014
Life Cycle Model

- Earnings
- Wealth
- Consumption
- Savings
- Dissaving

0: work starts  R: retirement  T: death
Rational vs. Myopic Individual

Rational individual
\([c_1 = c_1^*, c_2 = c_2^*]\)

Myopic individual
\([c_1 = W, c_2 = 0]\)
Adding forced savings $\tau=s^*$

Rational individual stays at $(c_1=c_1^*, c_2=c_2^*)$

Myopic individual moves to $(c_1=c_1^*, c_2=c_2^*)$
Figure 6. Possible effects of disability on prime-age male labour force participation

Source: Black, Furman, Rackstraw, Rao (2016)
**Figure 1.** Prime-age male labour force participation rate

Source: Black, Furman, Rackstraw, Rao (2016)
There was a pause in midlife mortality decline in the 1960s, largely explicable by historical patterns of smoking (13). Otherwise, the post-1999 episode in midlife mortality in the United States is both historically and geographically unique, at least since 1950. The turnaround is not a simple cohort effect; Americans born between 1945 and 1965 did not have particularly high mortality rates before midlife.

Fig. 2 presents the three causes of death that account for the mortality reversal among white non-Hispanics, namely suicide, drug and alcohol poisoning (accidental and intent undetermined), and chronic liver diseases and cirrhosis. All three increased year-on-year after 1998. Midlife increases in suicides and drug poisonings have been previously noted (14–16). However, that these upward trends were persistent and large enough to drive up all-cause midlife mortality has, to our knowledge, been overlooked. For context, Fig. 2 also presents mortality from lung cancer and diabetes. The obesity epidemic has (rightly) made diabetes a major concern for midlife Americans; yet, in recent history, death from diabetes has not been an increasing threat. Poisonings overtook lung cancer as a cause of death in 2011 in this age group; suicide appears poised to do so.

Table 1 shows changes in mortality rates from 1999 to 2013 for white non-Hispanic men and women ages 45–54 and, for comparison, changes for black non-Hispanics and for Hispanics. The table also presents changes in mortality rates for white non-Hispanics by three broad education groups: those with a high school degree or less (37% of this subpopulation over this period), those with some college, but no bachelor’s degree (31%), and those with a BA or more (32%). The fraction of 45- to 54-y-olds in the three education groups was stable over this period. Each cell shows the change in the mortality rate from 1999 to 2013, as well as its level (deaths per 100,000) in 2013.

Over the 15-y period, midlife all-cause mortality fell by more than 200 per 100,000 for black non-Hispanics, and by more than 60 per 100,000 for Hispanics. By contrast, white non-Hispanic mortality rose by 34 per 100,000. The ratio of black non-Hispanic to white non-Hispanic mortality rates for ages 45–54 fell from 2.09 in 1999 to 1.40 in 2013. CDC reports have highlighted the narrowing of the black–white gap in life expectancy (12). However, for ages 45–54, the narrowing of the mortality rate ratio in this period was largely driven by increased white mortality; if white non-Hispanic mortality had continued to decline at 1.8% per year, the ratio in 2013 would have been 1.97. The role played by changing white mortality rates in the narrowing of the black–white life expectancy gap (2003–2008) has been previously noted (17). It is far from clear that progress in black longevity should be benchmarked against US whites.

The change in all-cause mortality for white non-Hispanics 45–54 is largely accounted for by an increasing death rate from external causes, mostly increases in drug and alcohol poisonings and in suicide. (Patterns are similar for men and women when analyzed separately.) In contrast to earlier years, drug overdoses were not concentrated among minorities. In 1999, poisoning mortality for ages 45–54 was 10.2 per 100,000 higher for black non-Hispanics than white non-Hispanics; by 2013, poisoning mortality was 8.4 per 100,000 higher for whites. Death from cirrhosis and chronic liver diseases fell for blacks and rose for whites. After 2006, death rates from alcohol- and drug-induced causes for white non-Hispanics exceeded those for black non-Hispanics; in 2013, rates for white non-Hispanic exceeded those for black non-Hispanics by 19 per 100,000.

The three numbered rows of Table 1 show that the turnaround in mortality for white non-Hispanics was driven primarily by increasing death rates for those with a high school degree or less. All-cause mortality for this group increased by 134 per 100,000 between 1999 and 2013. Those with college education less than a BA saw little change in all-cause mortality over this period; those with a BA or more education saw death rates fall by 57 per 100,000. Although all three educational groups saw increases in mortality from suicide and poisonings, and an overall increase in external cause mortality, increases were largest for those with the least education. The mortality rate from poisonings rose more than fourfold for this group, from 13.7 to 58.0, and mortality from chronic liver diseases and cirrhosis rose by 50%. The final two rows of the table show increasing educational gradients from 1999 to 2013.

Fig. 2 presents the three causes of death that account for the mortality reversal among white non-Hispanics, namely suicide, drug and alcohol poisoning (accidental and intent undetermined), and chronic liver diseases and cirrhosis. All three increased year-on-year after 1998. Midlife increases in suicides and drug poisonings have been previously noted (14–16). However, that these upward trends were persistent and large enough to drive up all-cause midlife mortality has, to our knowledge, been overlooked. For context, Fig. 2 also presents mortality from lung cancer and diabetes. The obesity epidemic has (rightly) made diabetes a major concern for midlife Americans; yet, in recent history, death from diabetes has not been an increasing threat. Poisonings overtook lung cancer as a cause of death in 2011 in this age group; suicide appears poised to do so.

Table 1 shows changes in mortality rates from 1999 to 2013 for white non-Hispanic men and women ages 45–54 and, for comparison, changes for black non-Hispanics and for Hispanics. The table also presents changes in mortality rates for white non-Hispanics by three broad education groups: those with a high school degree or less (37% of this subpopulation over this period), those with some college, but no bachelor’s degree (31%), and those with a BA or more (32%). The fraction of 45- to 54-y-olds in the three education groups was stable over this period. Each cell shows the change in the mortality rate from 1999 to 2013, as well as its level (deaths per 100,000) in 2013.

Over the 15-y period, midlife all-cause mortality fell by more than 200 per 100,000 for black non-Hispanics, and by more than 60 per 100,000 for Hispanics. By contrast, white non-Hispanic mortality rose by 34 per 100,000. The ratio of black non-Hispanic to white non-Hispanic mortality rates for ages 45–54 fell from 2.09 in 1999 to 1.40 in 2013. CDC reports have highlighted the narrowing of the black–white gap in life expectancy (12). However, for ages 45–54, the narrowing of the mortality rate ratio in this period was largely driven by increased white mortality; if white non-Hispanic mortality had continued to decline at 1.8% per year, the ratio in 2013 would have been 1.97. The role played by changing white mortality rates in the narrowing of the black–white life expectancy gap (2003–2008) has been previously noted (17). It is far from clear that progress in black longevity should be benchmarked against US whites.

The change in all-cause mortality for white non-Hispanics 45–54 is largely accounted for by an increasing death rate from external causes, mostly increases in drug and alcohol poisonings and in suicide. (Patterns are similar for men and women when analyzed separately.) In contrast to earlier years, drug overdoses were not concentrated among minorities. In 1999, poisoning mortality for ages 45–54 was 10.2 per 100,000 higher for black non-Hispanics than white non-Hispanics; by 2013, poisoning mortality was 8.4 per 100,000 higher for whites. Death from cirrhosis and chronic liver diseases fell for blacks and rose for whites. After 2006, death rates from alcohol- and drug-induced causes for white non-Hispanics exceeded those for black non-Hispanics; in 2013, rates for white non-Hispanic exceeded those for black non-Hispanics by 19 per 100,000.

The three numbered rows of Table 1 show that the turnaround in mortality for white non-Hispanics was driven primarily by increasing death rates for those with a high school degree or less. All-cause mortality for this group increased by 134 per 100,000 between 1999 and 2013. Those with college education less than a BA saw little change in all-cause mortality over this period; those with a BA or more education saw death rates fall by 57 per 100,000. Although all three educational groups saw increases in mortality from suicide and poisonings, and an overall increase in external cause mortality, increases were largest for those with the least education. The mortality rate from poisonings rose more than fourfold for this group, from 13.7 to 58.0, and mortality from chronic liver diseases and cirrhosis rose by 50%. The final two rows of the table show increasing educational gradients from 1999 to 2013.
Figure 2.2: Employment of those aged 60–64

Source: Blundell, French, and Tetlow (2017)

Source: As Figure 2.1
Figure 2.2: Employment of those aged 60–64

Source: Blundell, French, and Tetlow (2017)
Figure 2.3: Employment of those aged 65–69

Source: Blundell, French, and Tetlow (2017)

Source: As Figure 2.1
Figure 2.3: Employment of those aged 65–69

Source: Blundell, French, and Tetlow (2017)

Source: As Figure 2.1
When the pension age was set at 65 in the UK, in 1925, life expectancy for men at that age was 11.2 years (as Figure 2.7 shows). This figure had changed little over the preceding 80 years. However, over the following 90 years (and particularly after 1960), it was to increase rapidly, reaching 18.9 years by 2012. This, coupled with the sharp fall in employment rates of older men described in section 2.2.1, led to a rapid expansion of the period spent in ‘retirement’.

The same coincidence of rising life expectancy and falling employment rates led to similar expansions in the prevalence and length of retirement across most developed countries after the Second World War. Most people in developed countries now expect to have a period of leisure at the end of their lives, with the date of their exit from employment determined not only by declining productivity and capacity to work but also by other factors such as their access to publicly and privately provided pensions.
The same eligibility age was adopted by the British, in 1909, when they too introduced an old age pension. For those who were reaching pension age in the UK system’s first year of operation, life expectancy at birth had been just 40 years for men and 43 years for women. Only one-in-four of those born in 1838 in the UK would actually have been alive to receive a pension.\footnote{In contrast, over four-in-five of the men born in 1943 and the women born in 1948 (who reached the eligibility age for public pensions in 2008) were still alive. Source: Department for Work and Pensions (2008).}

It was only somewhat later that pension eligibility ages were reduced to 65, which subsequently became widely accepted as an appropriate age to retire in many countries. The pension eligibility age was reduced to 65 in 1916 in Germany and in 1925 in the UK, and it was 65 from the inception of Social Security in 1935 in the US.\footnote{Age 65 had also been used by the Pensions Bureau in the US as the age of pension eligibility for Union army veterans from 1890 onwards (Costa 1998).}
effects cannot readily be separated.” Our paper helps to fill this gap, complementing a small set of papers that examine income effects in other disability contexts. Autor and Duggan (2007) and Autor et al. (2016) examine an income effect of changing access to Veterans’ Administration (VA) compensation for Vietnam War veterans on labor force participation, employment, and earnings. Marie and Vall Castello (2012) and Bruich (2014) study the income effect of DI benefits in Spain and Denmark, respectively. Finally, Deshpande (2016) studies the effect of children’s SSI payments on parents’ earnings. All of these studies find evidence consistent with substantial income effects in these other contexts. Our paper is the first to estimate an income effect specifically in the context of DI in the United States, which is the largest US federal expenditure on the disabled and one of the largest social insurance programs in the United States and around the world.

The remainder of the paper proceeds as follows. Section I describes the policy environment. Section II explains our identification strategy. Section III describes the data. Section IV shows our analysis of income effects. Section V discusses evidence on the extent to which income or substitution effects underlie earnings effects of DI by comparing our results to other literature. Section VI concludes. The online Appendix contains additional results.

5 Both studies estimate the reduced-form effects of receiving VA Disability Compensation. Autor et al. (2016, 3) conclude that “the effects that we estimate are unlikely to be driven solely by income effects.”

6 In the context of US Civil War veterans, Costa (1995) finds large income effects of pensions on labor supply.

7 Low and Pistaferri (2015) estimate many parameters simultaneously, including parameters of the work decision.
the finite-sample corrected Akaike Information Criterion (AICc). Using a baseline specification without additional controls, none of the specifications show that $\beta_2$ is statistically different from zero at the 5 percent level. Moreover, these regressions are rarely statistically significant for any polynomial order. The test that the coefficients are jointly significant across outcomes in the AICc-minimizing specifications shows $p = 0.20$ at the upper bend point and $p = 0.35$ at the lower.

We show in the online Appendix that there is no evidence for “bunching” in the density of initial AIME around the convex kink in the budget set created by the reduction in the marginal replacement rate around a bend point (since earning an extra dollar that increases AIME leads to a greater increase in DI benefits below the bend point than above it). Consistent with the exposition of the models in online Appendix 1, this finding could reflect that future DI claimants do not anticipate or understand the DI income they will receive, or that they do not react to the substitution incentives even when correctly anticipating them.

19 Working more will not lead to higher DI income if earnings are not in the highest earning years used to calculate AIME. However, as long as the prevalence of such cases evolves smoothly through the bend point (consistent with our data), the substitution effect should still lead to a greater incentive to earn below each bend point than above it.
to specification and misleading confidence intervals (which if anything applies still more in RKD settings; see Ganong and Jäger forthcoming, though note there is ongoing econometric discussion of this issue in Card et al. 2014).

Figure 5 shows the extensive margin, i.e., the fraction of the four years with positive annual earnings. There is an apparent increase in slope around the bend point. The regression analysis in Table 3 shows substantial effects in the linear specifications: a $1,000 increase in annual DI benefits is estimated to decrease the probability of reporting positive annual earnings by 1.29 percentage points in the specification without controls. As only a modest fraction of the sample has positive earnings in any given year, it makes sense that part of the observed earnings response would be operating through the extensive margin. Though these estimates remain positive under the quadratic and cubic specifications, they are smaller and lose statistical significance. In the online Appendix we also show similar patterns when the dependent variable is the probability of any employment over the full four years, rather than the percent of years with positive earnings (online Appendix Figure A6 and online Appendix Table A2). We conclude that there is some visual and statistical evidence of an employment effect at the upper bend point.

Notes: The figure shows the mean fraction of years when a beneficiary has positive annual earnings, over the four years after going on DI (i.e., the mean yearly employment rate over these four years), in $50 bins, as a function of distance from the bend point. The figure shows that the probability of positive earnings appears to slope upward more steeply above the upper bend point than below it.

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21 We obtain comparable results under specifications with the log odds of the employment rate as the dependent variable.

22 If DI benefits affect employment, then it is hard to interpret estimates of how DI payments affect earnings that are conditional on employment, as the sample is selected on an outcome (i.e., a beneficiary having positive earnings). The point estimates suggest insignificant negative impacts of DI benefits on earnings conditional on employment.
### Table 2—Smoothness of the Densities and Predetermined Covariates

<table>
<thead>
<tr>
<th>Dependent variable</th>
<th>Polynomial minimizing AICc</th>
<th>Estimated kink</th>
<th>Fraction of statistically significant kinks, polynomials of order 3–12</th>
</tr>
</thead>
<tbody>
<tr>
<td>Number of observations</td>
<td>9</td>
<td>−0.76 (1.41)</td>
<td>0%</td>
</tr>
<tr>
<td>Fraction male (× 1,000)</td>
<td>12</td>
<td>−0.100 (0.097)</td>
<td>0%</td>
</tr>
<tr>
<td>Average age when filing for DI (× 1,000)</td>
<td>10</td>
<td>1.27 (1.11)</td>
<td>40%</td>
</tr>
<tr>
<td>Fraction black (× 1,000)</td>
<td>12</td>
<td>−0.064 (0.048)</td>
<td>10%</td>
</tr>
<tr>
<td>Fraction of hearings allowances (× 1,000)</td>
<td>12</td>
<td>−0.024 (0.087)</td>
<td>0%</td>
</tr>
<tr>
<td>Fraction with mental disorders (× 1,000)</td>
<td>12</td>
<td>−0.075 (0.056)</td>
<td>10%</td>
</tr>
<tr>
<td>Fraction with musculoskeletal conditions (× 1,000)</td>
<td>12</td>
<td>0.081 (0.086)</td>
<td>0%</td>
</tr>
<tr>
<td>Fraction SSI recipients (removed from main sample) (× 1,000)</td>
<td>12</td>
<td>−0.034 (0.059)</td>
<td>0%</td>
</tr>
</tbody>
</table>

**Notes:** The table shows that the density of the assignment variable (i.e., initial AIME) and distributions of predetermined covariates are smooth around the upper bend point. We test for a change in slope at the bend point using polynomials of order 3 to 12. For each dependent variable, the table shows: the polynomial order that minimizes the corrected Akaike Information Criterion (AICc) (column 1), the estimated change in slope at the bend point and standard error under the AICc-minimizing polynomial order (column 2), and the percent of estimates of the change in slope that are statistically significant at the 5 percent level (column 3). Before running the regression, we take bin means of variables in bins of $50 width around the bend point, so each regression has 60 observations. See other notes to Table 1.

**Figure 4. Average Monthly Earnings after DI Allowance**

**Notes:** The figure shows mean monthly earnings in the first four years after going on DI, in $50 bins, as a function of distance of AIME from the bend point, where AIME is measured when applying for DI. The figure shows that mean earnings slope upward more steeply above the upper bend point than below it, with fitted lines that lie close to the data.
Figure 3. Annual Percent Mortality Rates around the Bend Points

A: Lower bend point

B: Family maximum bend point

C: Upper bend point

Notes: The figure shows the mean annual mortality rate in percent in the first four years after going on DI, in $50 bins, as a function of distance of AIME from the bend point. The figure shows that, at the lower and family maximum bend points, the mortality rate slopes upward more steeply above the bend point than below it, with fitted lines that lie close to the data.

Source: Gelber et al. (2018)
Figure 1: Job Exit Age Distribution (Full Sample)

Note: This figure shows the pooled distribution of job exit ages for all workers born between 1933 and 1948. The connected dots show the count of job exits within monthly bins. Vertical red lines indicate the location of main statutory ages throughout the sample period.

Data source: FDZ-RV - Themenfile SUFRZT1992-2014XVSBB_Siebold

Source: Siebold '17
Figure 2: Stylized Lifetime Budget Constraint

Note: The figure shows a stylized lifetime budget constraint for a worker who faces an Early Retirement Age of 60, a Full Retirement Age of 63 and an Normal Retirement Age of 65, who becomes eligible for a pathway requiring 35 years of contributions at age 58. The slope of the BC is the implicit net wage defined as $w_{net}^i = (1 - \tau_i)w_i$ as shown in section 2.3. The stylized shape of the constraint corresponds to incentives faced by the average worker: On average, workers face a 32% reduction in the implicit net wage (“kink size”) at age 60, a 42% reduction at age 63%, and a 21% increase in the implicit net wage at age 65.
Figure 4: Bunching at Specific Discontinuities

Panel A: Statutory age vs. pure financial incentive notch

Panel B: Statutory age vs. pure financial incentive kink

Note: This figure shows bunching at some cases of specific discontinuities. Panel titles indicate the type of discontinuity and panel subtitles indicate pathways and birth cohorts used. In panels A1, B1 and B2, the connected black dots show counts of job exit ages in monthly bins for the group indicated by the respective panel title. In panel A2, the black dots show counts of years of contributions instead. In all panels, the red line shows the counterfactual distribution estimated as a 7th-order polynomial, including round-age dummies in panels A1 and B1. Vertical red lines indicate the location of the discontinuity. $b$ is the excess mass, $d\tau/(1-\tau)$ is the change in the implicit net-of-tax rate at the discontinuity (kink size), and $\varepsilon$ is the implied elasticity of the retirement age w.r.t. the implicit net-of-tax rate. See appendix figure A2 for the lifetime budget constraints of the four groups.

Data source: FDZ-RV - Themenfile SUFRZ2N1992-2014XVSBB_Siebold
Figure 4: Bunching at Specific Discontinuities

Panel A: Statutory age vs. pure financial incentive notch

Panel B: Statutory age vs. pure financial incentive kink

Note: This figure shows bunching at some cases of specific discontinuities. Panel titles indicate the type of discontinuity and panel subtitles indicate pathways and birth cohorts used. In panels A1, B1 and B2, the connected black dots show counts of job exit ages in monthly bins for the group indicated by the respective panel title. In panel A2, the black dots show counts of years of contributions instead. In all panels, the red line shows the counterfactual distribution estimated as a 7th-order polynomial, including round-age dummies in panels A1 and B1. Vertical red lines indicate the location of the discontinuity. $b$ is the excess mass, $d\tau/(1-\tau)$ is the change in the implicit net-of-tax rate at the discontinuity (kink size), and $\varepsilon$ is the implied elasticity of the retirement age w.r.t. the implicit net-of-tax rate. See appendix figure A2 for the lifetime budget constraints of the four groups.

Data source: FDZ-RV - Themenfile SUFRTZN1992-2014XVSBB_Siebold
Figure 5: Bunching and Financial Incentives

Panel A: Statutory Ages

Panel B: Pure financial incentives

Note: The figure shows binned scatterplots of the excess mass at pure financial incentive discontinuities (panel A) and statutory ages (panel B) against kink size. In panel B, the type of statutory ages (Early, Full or Normal Retirement Age) controlled for. Each panel also includes the coefficient from a regression of normalized excess mass \( b/R \) on kink size, which can be interpreted as a difference-in-bunching elasticity, with bootstrapped standard error in parentheses. Appendix figure A3 shows additional graphs by type of statutory age.

Data sources: FDZ-RV - Themenfile SUFRTZN1992-2014XVSBB_Siebold
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Panel A: Statutory Ages

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Data sources: FDZ-RV - Themenfile SUFRTZN1992-2014XVSBB_Siebold

Source: Siebold '17
Panel B. Employment rates of men aged 60-64, 1970-2019

- **US** lowers early retirement age from 65 to 62 in 1961
- **Germany** lowers retirement age from 65 to 60 in 1973, increases it to 65 in 2000s

Source: Saez '21 using OECD database
Earning test for Social Security Benefit

- **Benefit**
- **Slope**: -0.5
- **Full benefit**
- **Phasing cut of benefit**
- **No benefit**

**Axes:**
- **Benefit**
- **Earnings**

**Values:**
- **$20k**
Americans making more money are living longer than those earning less. This means gaps in life expectancy by income have grown over time.

Average life expectancy at age 50

Position in women’s mid-career income distribution

Position in men’s mid-career income distribution

Source: Bosworth et al. 2016
Chart B—OASDI and HI Income and Cost as Percentages of Their Respective Taxable Payrolls

Chart A—OASI, DI, and HI Trust Fund Ratios
[Asset reserves as a percentage of annual cost]
Automatic enrollment effect

Automatic enrollment dramatically increases participation.

401(k) participation by tenure at firm: Company B

Source: Madrian and Shea (2001)
Automatic enrollment effect

Employees enrolled under automatic enrollment cluster at the default contribution rate.

Distribution of contribution rates: Company B

Source: Madrian and Shea (2001)
The Flypaper Effect in Individual Investor Asset Allocation (Choi, Laibson, Madrian 2007)

Studied a firm that used several different match systems in their 401(k) plan. I’ll discuss two of those regimes today:

Match allocated to employer stock and workers can reallocate
  • Call this “default” case (default is employer stock)

Match allocated to an asset actively chosen by workers; workers required to make an active designation.
  • Call this “no default” case (workers must choose)

Economically, these two systems are identical. They both allow workers to do whatever the worker wants.

Source: courtesy of David Laibson
Consequences of the two regimes

<table>
<thead>
<tr>
<th></th>
<th>Default</th>
<th>No Default</th>
</tr>
</thead>
<tbody>
<tr>
<td>Own Balance in Employer Stock</td>
<td>24%</td>
<td>20%</td>
</tr>
<tr>
<td>Matching Balance in Employer Stock</td>
<td>94%</td>
<td>27%</td>
</tr>
<tr>
<td>Total Balance in Employer Stock</td>
<td>56%</td>
<td>22%</td>
</tr>
</tbody>
</table>

Source: courtesy of David Laibson
Active decision effect on participation

401(k) participation increases substantially when employees are not allowed to be passive about savings.

Source: courtesy of David Laibson
Employer match threshold and contribution rates

Changing the match threshold caused employees to slowly move from the old threshold to the new threshold.

401(k) contribution rate response to match threshold change: Company G

Source: courtesy of David Laibson