Pension Access and Black Men's Mental Health: Evidence from South Africa's 2008–2010 Policy Reform

Ugo Oyefia

Department of Sociology, University of Kansas

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Correspondence: Ugo Oyefia, ugo@ku.edu.
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Abstract

South Africa's pension reform brought men's qualifying age into line with women's, extending eligibility for income support earlier in life. I use longitudinal data from the National Income Dynamics Study to examine how pension access relates to health among older Black African men. At the qualifying-age threshold, men report greater pension receipt and higher income from adult benefits. The health estimates show no consistent improvement in depressive symptoms or self-rated health, and narrower age comparisons suggest worse mental health. Transfers from disability grants also mean that a newly reported pension need not represent entirely new support. The findings show why greater pension receipt alone is insufficient evidence of a health gain. Because the comparisons pool old and new age rules, I cannot isolate the reform's health consequences.

Keywords: social protection; mental health; aging; South Africa; pensions

Introduction

South Africa lowered men's qualifying age for the Old Age Grant from 65 to 60 between 2008 and 2010, matching women's age of qualification (Republic of South Africa 2008). Older Black African men approached this reform after decades of racial restrictions on land and employment. In his account of Kas Maine's life, van Onselen (1996) shows how these restrictions constrained one family's efforts to make a living. A regular pension could ease the strain of meeting daily expenses in later life, making earlier eligibility relevant to mental health as well as income (Pearlin et al. 1981).

This paper examines the relationship between pension access and health among older Black African men using the National Income Dynamics Study (NIDS). I first assess whether men report greater pension receipt and additional income from public benefits at the qualifying-age threshold, then examine depressive symptoms and self-rated health at the same threshold. Separating these questions matters because a rise in pension receipt need not represent the same financial change for every recipient. The comparisons use the rules in place at each interview and include observations before and after the reform; they assess health around age qualification, not the effect of lowering the qualifying age itself.

Background

The Old Age Grant provides means-tested social assistance. Reaching the qualifying age did not establish full eligibility: applicants also had to meet residence and institutional-care requirements, and the means test assessed income and assets, including spousal resources (Department of Social Development 2008). My analysis identifies age qualification rather than whether men met every condition. The Appendix details the historical rules and their implementation.

The early rollout converted more than 36,000 disability grants for men aged 63–64 into old-age pensions (Social Security Administration 2008). These men were already receiving public assistance, unlike men entering the system for the first time. I therefore examine pension receipt alongside broader measures of benefit access and income.

The stress-process perspective offers a reason to expect mental health benefits when a pension improves financial security. Persistent difficulties in meeting material needs can generate psychological distress, while access to resources can reduce that strain (Pearlin et al. 1981). A regular payment could help with recurring expenses and reduce uncertainty about the income available to meet them. If pension access reduces these pressures, I would expect fewer depressive symptoms among men who have reached the qualifying age. Pension studies

in China and Mexico report mental health gains, and a broader review of cash transfers finds small average improvements in mental health and subjective well-being (Chen, Wang, and Busch 2019; Galiani, Gertler, and Bando 2016; McGuire, Kaiser, and Bach-Mortensen 2022).

South African evidence places these financial changes within households rather than treating the pension as a resource that only its recipient uses. Case (2004) links pension income to health in households that pool income. The nutritional gains children receive can differ according to the pension recipient's gender (Duflo 2003), and pension income can also change living arrangements (Hamoudi and Thomas 2014). These findings matter for evaluating men's pensions: shared resources may affect a man's health even when he is not the recipient, while a pension he receives may benefit other household members. His reported grant and his own health consequently capture only part of the pension's potential consequences.

Any health benefit may also depend on how long men have received a pension, since relief from current expenses need not undo disadvantages accumulated over a lifetime (Pearlin et al. 2005). South African studies document gendered patterns in well-being around pension eligibility and associations between longer pension exposure and better health, particularly physical disability outcomes (Riumallo Herl et al. 2022; Schatz et al. 2012). The distinction is relevant here because men's ages at interview indicate whether they have reached the qualifying age, but not how long they have received payments.

Data and Methods

I follow 523 Black African men aged 55–69 at baseline through NIDS Waves 1–3, covering 2008–2012 (Southern Africa Labour and Development Research Unit 2018a; 2018b; 2018c). I do not add younger entrants; the age window and nonmissing outcomes determine each estimation sample. Table 1 reports descriptive statistics, and the Appendix details sample retention.

My principal outcome is the ten-item Center for Epidemiologic Studies Depression scale (CES-D-10), which ranges from zero to 30, with higher scores indicating more symptoms. I require at least eight completed items and also examine scores of ten or higher as screen positivity, not a clinical diagnosis (Andresen et al. 1994). Because optimal cutoffs differ across South African language samples, I emphasize the continuous score (Baron, Davies, and Lund 2017).

I code self-rated health as one for excellent, very good, or good health and zero for fair or poor health. I measure state-pension and disability-grant receipt separately, then combine

both with war-veteran benefits to measure receipt of any adult benefit and the total monthly amount in nominal rand. I retain disability recipients in the main sample.

I estimate regression discontinuity models within eight years of qualification, using age in months and separate linear age trends on either side. Models include wave-by-qualifying-age fixed effects and survey/panel weights; I cluster standard errors by baseline primary sampling unit. The Appendix specifies the timing rules and estimation details.

Interpreting the threshold difference as a health effect requires other influences on health to change smoothly across qualification and the modeled age trends to be adequate (Hahn, Todd, and van der Klaauw 2001; Lee and Lemieux 2010). These models pool observations under old and new age rules rather than estimating within-person changes. Secondary two-stage least squares models use age qualification to instrument pension receipt; the Appendix explains their additional assumptions and reports the diagnostic tests.

Results

Men within the main age window average about 62 years of age at baseline (Table 1). Pension receipt rises and average depressive symptoms decline across waves. These trends do not isolate a pension effect because the cohort also ages and its composition changes.

Pension receipt is 31.42 percentage points higher at age qualification (Table 2, Panel A). The increase in receipt of any adult benefit is smaller, as I would expect if some men moved between programs. Nevertheless, combined monthly income from these benefits is approximately R301 higher above the threshold. The income difference suggests that greater pension receipt is not simply a change in benefit labels, although the uncertain disability estimate and differing outcome samples prevent a precise accounting of grant conversion.

Receipt below the qualifying age reflects survey reports, not verified eligibility. Some cases depend on the 2008 rollout date; others remain unexplained and may reflect age or benefit-reporting errors.

For depressive symptoms, the estimated difference at qualification is 0.47 points on the CES-D-10 (95 percent confidence interval [CI]: $[-1.10, 2.03]$). Higher scores indicate worse mental health, so the point estimate does not favor men above the threshold. Its interval is too wide to distinguish improvement from deterioration. The descriptive age pattern also varies substantially on both sides of qualification (Appendix Figure A2).

The models estimate lower probabilities of both screen positivity and good health above the qualifying age. The point estimates thus suggest opposing health patterns, but both intervals include zero. None of the health-outcome tests rejects after Holm adjustment. The secondary pension-receipt estimates have still wider intervals (Table 2, Panel B).

Table 1

Sample Characteristics by Survey Wave

Characteristic	2008		2010–2011		2012	
	M or %	*n*	*M* or %	*n*	*M* or %	*n*
Age in 2008 (years)	62.07	425	59.20	288	58.64	235
	(0.29)		(0.24)		(0.22)	
School-grade equivalent	4.45	424	5.05	284	5.26	232
	(0.32)		(0.34)		(0.35)	
Partnered (%)	79.62	424	73.26	286	75.66	234
	(2.83)		(4.17)		(4.48)	
Urban residence (%)	48.24	425	51.86	288	57.02	235
	(5.07)		(6.34)		(6.67)	
State pension (%)	27.87	425	44.90	286	61.58	235
	(2.93)		(3.77)		(5.37)	
CES-D-10 score	8.55	425	7.75	276	7.45	234
	(0.27)		(0.32)		(0.30)	
Screen positive (%)	35.38	425	25.75	276	22.66	234
	(3.35)		(3.69)		(3.02)	
Good self-rated health (%)	59.99	422	78.92	286	73.89	235
	(4.13)		(3.49)		(4.67)	

Note. M = weighted mean; percentages use survey/panel weights; n = nonmissing observations within eight years of qualification. Parentheses contain standard errors, which I cluster by baseline primary sampling unit. The first four rows describe baseline characteristics. School-grade equivalents range from 0 to 12; no schooling is 0 and National Technical Certificate levels 1–3 are 10–12; other codes are missing. CES-D = Center for Epidemiologic Studies Depression scale. Screen positive denotes a score of 10 or higher. Good health includes excellent, very good, and good responses. Data are from the National Income Dynamics Study, Waves 1–3.

Figure 1 shows how the depressive-symptom estimates change across age windows. The narrower comparisons suggest worse mental health rather than a stable improvement. Within three years of qualification, the symptom estimate rises to 2.71 points (95 percent CI: $[0.39, 5.03]$), and screen positivity is also higher with an interval excluding zero. I do not adjust tests across the full sensitivity-analysis set, and the corresponding pension-receipt estimates remain inconclusive. Appendix Table A2 reports the estimates and sample sizes; the Appendix also describes the remaining sensitivity checks.

Table 2

Age Qualification, Pension Receipt, and Health

Outcome	b	SE	95% CI	p	n
Panel A. Differences at age qualification					
State pension	31.42	7.68	[16.29, 46.56]	< .001	946
Disability grant	−8.36	4.37	[−16.96, 0.24]	.057	946
Any adult benefit	23.01	7.59	[8.06, 37.96]	.003	947
Combined amount (rand)	300.75	111.76	[80.58, 520.92]	.008	942
CES-D-10 score	0.47	0.79	[−1.10, 2.03]	.558	935
Screen positive	−3.60	8.21	[−19.78, 12.57]	.661	935
Good self-rated health	−6.93	7.05	[−20.82, 6.95]	.326	943
Panel B. Pension receipt: two-stage least squares					
CES-D-10 score	1.51	2.69	[−3.76, 6.78]	.574	934
Screen positive	−11.83	26.30	[−63.38, 39.72]	.653	934
Good self-rated health	−22.20	23.72	[−68.68, 24.28]	.349	942

Note. b = unstandardized coefficient; SE = standard error; CI = conventional confidence interval; n = person-wave observations. Panel A uses available observations for each outcome. Panel B instruments pension receipt with age qualification on outcome-matched samples (first stages in Appendix Table A1). Models use an eight-year window, separate age slopes, wave-by-cutoff fixed effects, and survey/panel weights; I cluster standard errors by baseline primary sampling unit. CES-D = Center for Epidemiologic Studies Depression scale. Binary coefficients are percentage points; CES-D coefficients are scale points; amounts are nominal rand. Screen positive denotes scores of 10 or higher. Good health includes excellent, very good, and good responses. Adult benefits comprise state pensions, disability grants, and war-veteran benefits. The p values are unadjusted. Data are from the National Income Dynamics Study, Waves 1–3.

Discussion

At age qualification, older Black African men report greater pension receipt and higher income from adult benefits, but the health findings do not follow the same pattern. The main estimates are compatible with both health gains and losses, and narrower comparisons suggest worse mental health. These results do not reproduce the mental health improvements that pension studies in China and Mexico report (Chen et al. 2019; Galiani et al. 2016), although the studies use different designs. They also do not provide a stable basis for concluding that pensions have no health consequences.

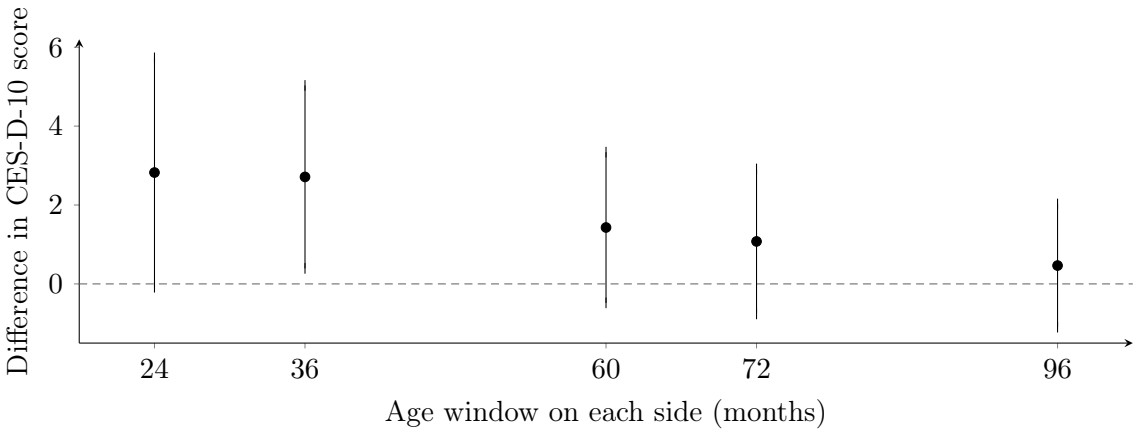

Figure 1. Estimated Differences in Depressive Symptoms by Age Window

Source: National Income Dynamics Study, Waves 1–3. *Note:* Points show estimated differences at age qualification; bars are conventional baseline-PSU-clustered 95 percent intervals. The main model uses 96 months on each side. All windows use separate age slopes and the main timing rule. Positive values indicate more symptoms above the qualifying age.

A pension can represent different financial changes for men entering public assistance and those transferring from disability grants. Higher combined benefit income indicates more than a change in reported pension status, but the estimates do not separate these experiences. The stress-process argument concerns relief from financial strain, which benefit amounts alone cannot establish. I do not observe whether men became better able to meet expenses or how household members shared their income.

Following a fixed cohort leaves few men below the qualifying age in the final wave as the same men grow older and some no longer complete interviews. Weights cannot supply the missing age support. The broad main window relies heavily on modeled age trends; the different findings in narrower windows show how consequential the choice of comparison is. Birth-month error and the small initial age-63 group add uncertainty. I lack complete means-test information and first-payment dates, and the pooled old and new age rules do not isolate the reform's effect. These limitations preclude attributing the adverse narrower-window findings to the policy or interpreting the broad-window estimates as evidence that earlier payments were ineffective.

An assessment of the reform must also consider benefits that an individual health measure would miss. South African research shows that pensions can benefit other household members, making a man's symptoms an incomplete measure of their value (Case 2004; Duflo 2003; Hamoudi and Thomas 2014). Whether earlier eligibility improved men's lives therefore requires evidence about financial security and health after sustained receipt.

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Appendix

Data and policy details

The 2008 amendment specified a male qualifying age of 63 after April 1, 2008, 61 after April 1, 2009, and 60 after April 1, 2010. Although the first statutory date was in April, the Act commenced on July 14, 2008, and the ministry subsequently described payments to newly age-qualified men as beginning in July (Molewa 2009; Republic of South Africa 2008). The August 2008 regulations set residence and institutional-care requirements and a means test covering income and assets, including spousal resources for married applicants (Department of Social Development 2008). Income and asset thresholds later changed, and a 2012 amendment explicitly included refugees in the relevant status provision (Department of Social Development 2012; National Treasury 2011). I do not use personal income alone to approximate the historical means test.

The main timing rule retains age 65 through June 2008, omits July because implementation began within that month, and uses age 63 from August 2008 through March 2009, age 61 through March 2010, and age 60 thereafter. An alternative rule follows the April 2008 schedule. I hold each man's baseline birth report fixed across waves to prevent later revisions from creating artificial threshold crossings. These rules identify age qualification at interview, not the date of application or first payment.

I use the adult and derived files from NIDS Wave 1 (2008), Wave 2 (2010–2011), and Wave 3 (2012). The baseline cohort includes completed interviews with valid birth and interview dates. Of the 523 baseline men, 378 completed Wave 2 interviews and 349 completed Wave 3 interviews. I calculate age in months from each interview's month and year and the birth month and year each man reported at baseline.

I code CES-D items from zero to three and reverse the two positively worded items. For men answering at least eight items, I multiply the item mean by ten. A sensitivity analysis instead requires all ten items. I code self-rated health as one for excellent, very good, or good health and zero for fair or poor health. Adult benefits comprise state pensions, disability grants, and war-veteran benefits; combined amounts are monthly nominal rand, not total household income or consumption.

Figures A1 and A2 show the descriptive age patterns around qualification.

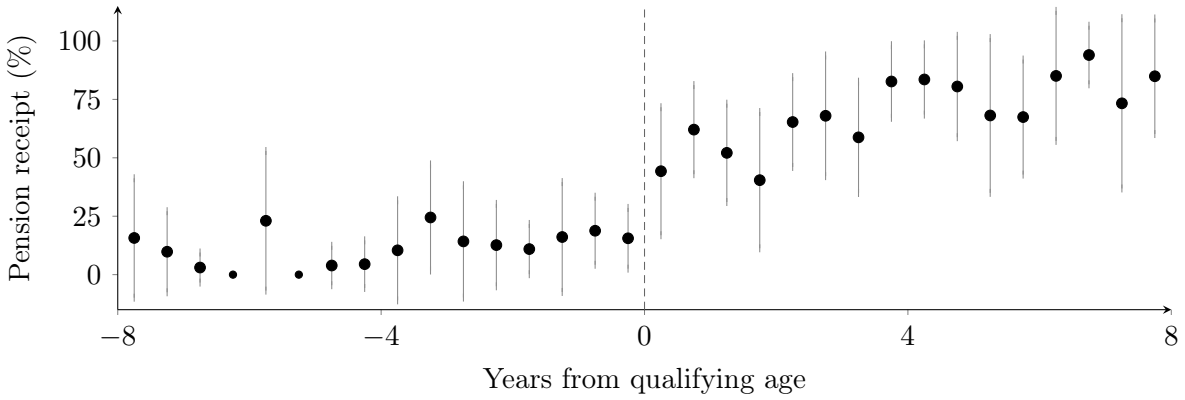

Figure A1. Reported Pension Receipt by Age Relative to Qualification

Source: National Income Dynamics Study, Waves 1–3. *Note:* Points are survey/panel-weighted means in six-month bins; bars are baseline-PSU-clustered 95 percent intervals where available. Two all-zero bins have no reported interval. The dashed line marks the threshold. Bins are descriptive and do not adjust for wave-by-cutoff composition.

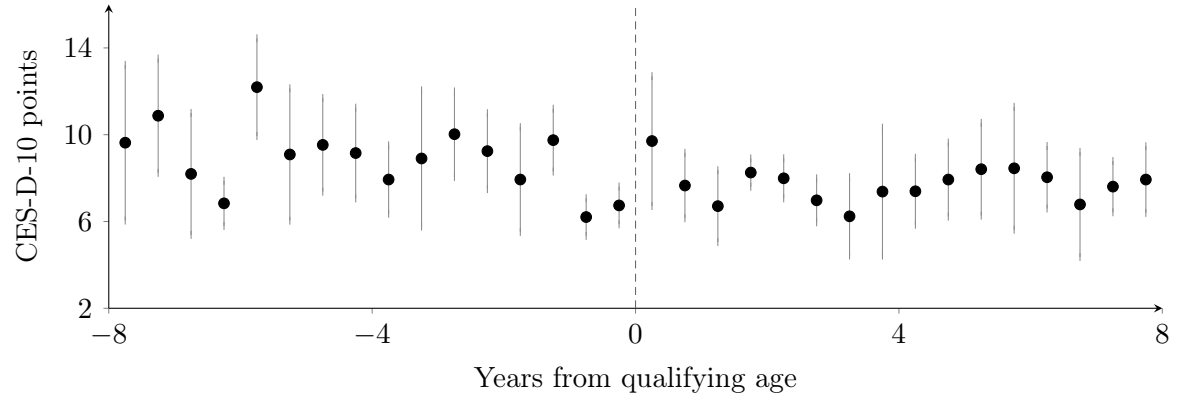

Figure A2. Depressive Symptoms by Age Relative to Qualification

Source: National Income Dynamics Study, Waves 1–3. *Note:* Higher scores indicate more symptoms. Points and intervals follow Figure A1. I require at least eight completed items and prorate scores to ten.

Model specification and age support

Let x_{it} denote months from the qualifying age, $Z_{it} = 1[x_{it} \geq 0]$, and c index the wave-by-cutoff cell. The main age-qualification model is

$$Y_{it} = \alpha_c + \tau Z_{it} + \beta x_{it} + \gamma x_{it} Z_{it} + \varepsilon_{it}. \tag{A1}$$

The model uses a common threshold difference and common side-specific age slopes across cells, with no additional individual covariates. The main window is 96 months on either

side, with uniform weighting over age. Descriptive statistics and models use baseline survey weights or the relevant later-wave panel weights. I cluster standard errors by baseline primary sampling unit (PSU), keeping each man's repeated observations together. I report conventional cluster-based intervals, not regression-discontinuity-bias-corrected intervals, and express binary-outcome estimates in percentage points.

For CES-D, the Wave 1 age-65 cell contains 263 observations below the threshold and 133 at or above it; the age-63 cell contains 23 and six. The corresponding counts are 82 and 194 in Wave 2 and 30 and 204 in Wave 3. The sample contains no age-61 cell. Because I follow the 2008 cohort without adding younger entrants, Wave 3 contains only about one year of below-threshold age support.

Instrumental-variable (IV) models use age qualification to instrument pension receipt. Each first stage, reduced form, and IV estimate uses the same outcome-specific sample. Interpreting an IV estimate as a local receipt effect requires relevance, exclusion, and monotonicity in addition to continuity (Hahn et al. 2001). Grant conversion complicates the treatment interpretation because receipt does not represent a fixed increment in income. I report conventional intervals and clustered Anderson–Rubin confidence sets (Anderson and Rubin 1949); the latter address weak-instrument uncertainty under instrument validity, not bias from an incorrect age trend.

Table A1 reports matched-sample first stages and Anderson–Rubin sets for the receipt estimates in Table 2, Panel B. Panel A uses all available observations for each outcome, so its pension-receipt row is not the matched first stage for Panel B.

Table A1

Matched First Stages and Confidence Sets for Pension-Receipt Estimates

Health outcome	First-stage b	F	95% AR Set	n
CES-D-10 score	30.68	15.53	$[-3.40, 9.10]$	934
Screen positive	30.68	15.53	$[-70.00, 50.00]$	934
Good self-rated health	31.10	16.39	$[-86.50, 22.50]$	942

Note. First-stage b = pension-receipt difference at age qualification, in percentage points; F = first-stage test statistic; AR = Anderson–Rubin confidence set for the health effect of pension receipt; n = person-wave observations. AR sets use CES-D-10 scale points or percentage points for binary health outcomes and assume instrument validity. Symptom and screening models include 507 men; self-rated health includes 505. Models match the specification in Table 2, Panel B, and cluster by 240 baseline primary sampling units. Data are from the National Income Dynamics Study, Waves 1–3.

I obtain Anderson–Rubin sets by inverting tests over grids from −30 to 30 in .05-point increments for CES-D and from −2 to 2 in .005 increments for binary outcomes before percentage-point conversion. Each accepted set is contiguous and does not reach a grid boundary; endpoints are approximate.

I report unadjusted and Holm-adjusted p values for the primary health outcomes (Holm 1979). The adjusted values are 1.000 for CES-D, 1.000 for screen positivity, and .979 for good health. Table 2 reports unadjusted two-sided tests.

Alternative comparisons and diagnostics

Table A2 supplies the coefficients and sample sizes plotted in main-text Figure 1. In the 36-month window, the screen-positivity estimate is 27.55 percentage points ($p = .020$); the corresponding receipt-IV tests do not reject at five percent.

Table A2

Depressive-Symptom Estimates by Symmetric Age Window

Window (months)	b	SE	95% CI	p	n
24	2.82	1.47	$[-0.08, 5.73]$	.057	323
36	2.71	1.18	$[0.39, 5.03]$	.022	465
60	1.43	0.97	$[-0.48, 3.34]$	.142	691
72	1.08	0.93	$[-0.76, 2.91]$	.249	768
96	0.47	0.79	$[-1.10, 2.03]$	.558	935

Note. b = unstandardized difference at age qualification; SE = standard error; CI = confidence interval; n = person-wave observations. Coefficients are Center for Epidemiologic Studies Depression scale points. Windows extend the stated number of months on each side of the qualifying age. Models use separate age slopes, wave-by-cutoff fixed effects, and survey/panel weights. I cluster standard errors by baseline primary sampling unit. I do not adjust two-sided p values across sensitivity tests. Data are from the National Income Dynamics Study, Waves 1–3.

Applying the statutory April 2008 schedule yields 0.02 CES-D points; imposing the alternative personal-income restriction yields −0.62 points; excluding six months on either side of the threshold yields −0.70 points. All three intervals include zero. Restricting the analysis to Waves 1–2 yields 0.03 points (95 percent CI: $[-1.33, 1.38]$).

Cell-specific models require at least 15 observations and three distinct monthly age values on each side. The Wave 1 age-63 cell fails this requirement. Other cells use local linear fits,

quadratic bias correction, triangular kernels, mean-squared-error-selected bandwidths, mass-point adjustment, and baseline-PSU clustering (Calonico, Cattaneo, and Titiunik 2014). Because monthly age is discrete, these checks remain sensitive to the available age support (Kolesár and Rothe 2018). All cell-specific reduced-form intervals include zero. The Wave 3 fuzzy CES-D estimate is 23.39 points (95 percent CI [0.27, 46.50]; $p = .047$). The wide conventional fuzzy Wald interval and asymmetric age support warrant caution, especially because this is one result among many checks.

Additional specifications vary age slopes and kernels; add baseline education, partnership, and urban residence; use alternative survey weights or respondent clustering; vary birth-date and interview-date coding; omit the birthday month or six months around the cutoff; restrict observations to Waves 1–2 or a balanced panel; require complete CES-D items; and alter the personal-income screen or overlapping disability definition. Baseline covariates may already reflect earlier pension access, so their inclusion is a specification check rather than a separate identification strategy.

No balance, missingness, placebo, or retention test rejects after its within-family Holm correction. Non-rejection does not establish continuity or eliminate possible selection bias. Holm corrections cover the three primary health outcomes and separate diagnostic families, not the entire sensitivity-analysis set.

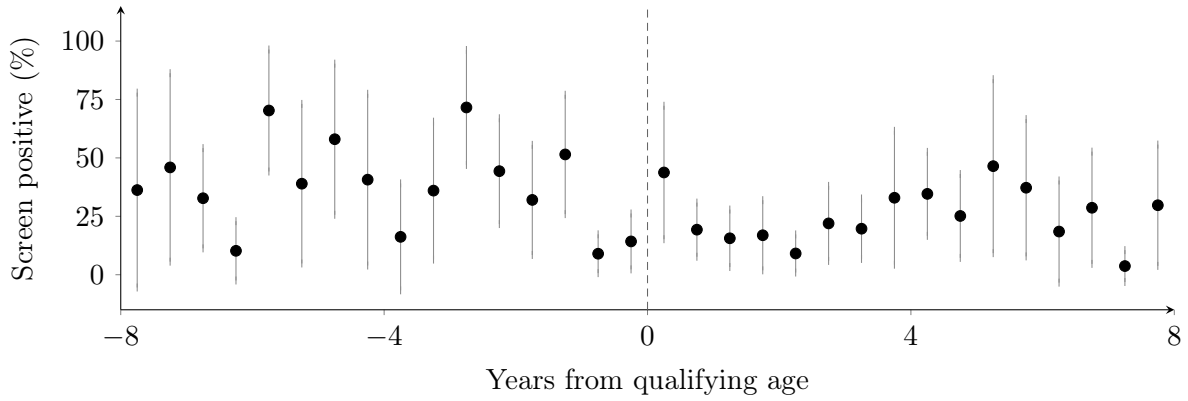

Figure A3. CES-D-10 Screen Positivity by Age Relative to Qualification

Source: National Income Dynamics Study, Waves 1–3. *Note:* A score of ten or more defines the screen. Points and intervals follow Figure A1. I do not constrain linear mean intervals to 0–100 percent; sparse-bin coverage may be poor.

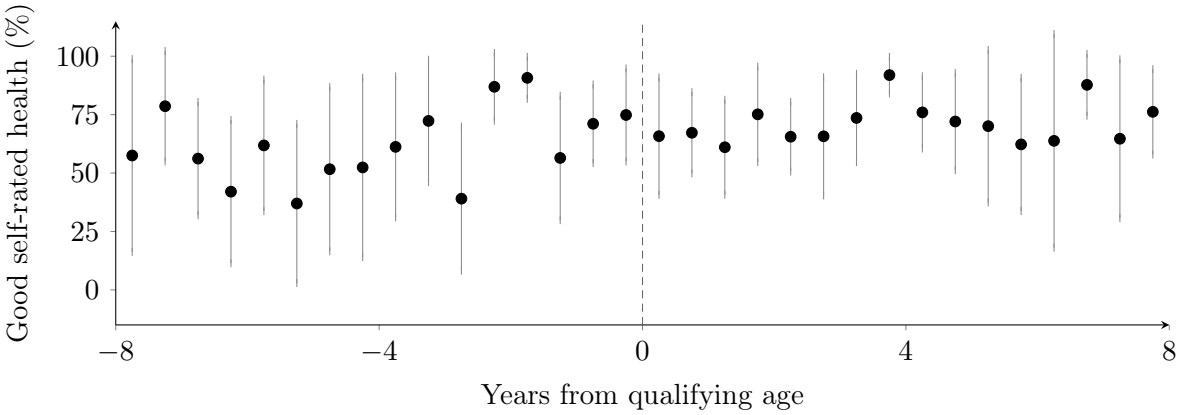

Figure A4. Good Self-Rated Health by Age Relative to Qualification

Source: National Income Dynamics Study, Waves 1–3. *Note:* Good health includes excellent, very good, and good responses. Points and intervals follow Figure A1. I do not adjust bins for wave composition; sparse-bin intervals warrant caution.