Research Square

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Blood pressure control transitions and subsequent reported stroke among Chinese adults with hypertension in a landmark cohort study

Binbin Zhang
the Eighth Affiliated Hospital of Sun Yat- sen University
Shafiu A Umar Shinge
the Eighth Affiliated Hospital of Sun Yat- sen University
Baixin Zheng
the Eighth Affiliated Hospital of Sun Yat- sen University
Yongjia Qiang
the Eighth Affiliated Hospital of Sun Yat- sen University
Zhuxuan Liu
the Eighth Affiliated Hospital of Sun Yat- sen University
Lu Zhang
the Eighth Affiliated Hospital of Sun Yat- sen University
Minnan Gao
the Eighth Affiliated Hospital of Sun Yat- sen University
Bin Li
First Affiliated Hospital of Kunming Medical University
Kuan Zeng
the Eighth Affiliated Hospital of Sun Yat- sen University
Yanqi Yang
yangyq@mail.sysu.edu.cn

the Eighth Affiliated Hospital of Sun Yat- sen University

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Abstract

Background

Blood pressure measured at successive visits may identify different subsequent stroke risks, but changes in measured control should be distinguished from treatment effects. We examined blood pressure control transitions from 2011 to 2013 in the China Health and Retirement Longitudinal Study (CHARLS).

Methods

This landmark analysis included adults aged at least 45 years with baseline hypertension, valid blood pressure measurements at both visits, and no reported stroke before the landmark. Control was defined as systolic pressure below 140 mmHg and diastolic pressure below 90 mmHg. Subsequent first reported physician-diagnosed stroke through the 2020 survey was analysed with an interval-censored Weibull model and community-clustered uncertainty. The primary comparison was uncontrolled-to-controlled versus uncontrolled-to-uncontrolled pressure among initially uncontrolled participants. Covariates included baseline demographic and clinical characteristics and continuous blood pressure.

Results

The complete-case stroke sample comprised 2902 participants with 256 events. Among 2126 initially uncontrolled participants, 190 events occurred. In the primary Weibull model, transition to control was associated with a lower modelled stroke hazard than uncontrolled pressure at both visits (hazard ratio [HR], 0.68; 95% confidence interval [CI], 0.49–0.95). Multiple imputation gave an HR of 0.70 (95% CI, 0.51–0.97). A semiparametric interval-censored model gave an HR of 0.69 (community-bootstrap 95% interval, 0.46–0.96). Several exposure-definition analyses could not exclude no association. A lognormal model had a lower Akaike information criterion than the Weibull model and yielded an uncertain time ratio of 1.22 (95% CI, 0.97–1.53).

Conclusions

In the primary Weibull model, transition to measured control was associated with a lower modelled hazard of subsequently reported stroke. The evidence depended on distributional assumptions and exposure definitions. Death-related underascertainment, measurement error, and selection limit interpretation; these estimates do not establish a treatment effect.

Clinical trial registration

Clinical trial number: not applicable. This study was an observational secondary analysis of existing CHARLS cohort data and did not involve the assignment of any healthcare intervention to participants.

Background

Blood pressure lowering reduces cardiovascular events in randomized evidence, including stroke [1]. In observational cohorts, however, a lower pressure at a later visit can reflect several processes: treatment, behavioural change, intercurrent illness, measurement variability, or regression to the mean. Understanding the prognosis associated with a change in measured control therefore requires a design that separates the exposure assessment period from subsequent outcome observation.

Large observational studies relate usual blood pressure to vascular mortality [2]. Randomized evidence includes intensive-control strategies in SPRINT, older Chinese adults in STEP, people with type 2 diabetes in ACCORD, and secondary stroke prevention in PROGRESS [3, 4, 5, 6]. These trials address distinct populations and interventions, while participant-level meta-analysis evaluates pharmacological lowering across baseline pressure levels [7]. Their causal questions differ from the prognosis associated with two observed control states. Large Chinese surveys also document gaps in hypertension awareness, treatment, and control [8, 9], motivating investigation of longitudinal control patterns without equating measured attainment with assigned treatment.

CHARLS provides repeated health interviews and blood pressure measurements in middle-aged and older Chinese adults [10]. Related analyses have examined cumulative blood pressure and cardiovascular outcomes [11], hypertension care-cascade categories and cardiovascular disease [12], and factors associated with transitions in blood pressure status [13]. These studies establish substantial overlap with the present topic. The narrower question addressed here is whether attaining measured control at a second visit is associated with subsequent reported stroke among participants whose pressure was initially uncontrolled. A complementary analysis examines loss of control among initially controlled participants.

We used a 2013 landmark after blood pressure assessment in 2011 and 2013 and reconstructed stroke histories from contemporaneous questionnaires. Because an affirmative response can represent a correction to earlier disease history, and event dates are incompletely observed, we explicitly examined historical corrections, interval censoring, and alternative outcome and exposure definitions. The objective was to estimate an observational association, rather than a treatment effect or an optimal blood pressure target.

Methods

Study design and participants

CHARLS is a longitudinal study of Chinese adults aged 45 years and older and their spouses [10, 14]. Reporting follows STROBE, with a section-mapped checklist in Additional file 1 [15, 16]. We linked the

local Harmonized CHARLS version D release with original 2011, 2013, 2015, 2018, and 2020 questionnaire files, blood pressure records, sample information, and the 2020 exit module. The analysis was restricted to participants interviewed in 2011 who were at least 45 years old. Baseline hypertension was defined by measured systolic pressure of at least 140 mmHg or diastolic pressure of at least 90 mmHg, contemporaneously reported hypertension, or reported use of western antihypertensive medication.

Participants needed valid pressure measurements and explicit negative stroke histories in 2011 and 2013. Of 3250 meeting these initial criteria, 113 were excluded after later responses indicated a prior or temporally ambiguous stroke history, leaving 3137. We excluded 154 without analysable post-landmark stroke observation time and one with a chronology conflict between interview and death information. Among the remaining 2982, 80 had missing clinical covariates; the primary complete-case sample therefore included 2902 participants (Fig. 1). Eligibility at a later landmark necessarily conditions on survival and attendance during the exposure window.

Blood pressure and transition groups

We audited paired systolic and diastolic readings against the original biomarker files. The 2013 biomarker questionnaire specifies an Omron HEM-7200 monitor, three measurements, a seated and quiet participant, and an arm supported at heart height; the introductory protocol uses the left arm, while actual arm, posture, compliance, and recent smoking, exercise, alcohol or food are recorded separately [17]. Interviewers are instructed to demonstrate the procedure. These source-specific details are distinguished from general blood pressure measurement recommendations [18]. Cuff dimensions, a fixed premeasurement rest duration, and adherence to a specified interreading interval were not verified in the available participant records and are not assumed. Interview-month distributions and recorded measurement conditions are reported in Additional file 1. Valid pairs required systolic pressure of 70–260 mmHg, diastolic pressure of 40–150 mmHg, and systolic pressure greater than diastolic pressure. At least two valid pairs were required at each visit. The primary visit-level measure was the mean of the last two valid pairs. A sensitivity analysis used the mean of all valid pairs. Using later valid readings was an analytical choice intended to reduce sensitivity to an initial reading, not a claim that this averaging rule was mandated by CHARLS. Post-revision analyses additionally restricted the sample to seated, fully compliant measurements at both visits and examined continuous systolic and diastolic decreases. None of these analyses calibrates usual pressure or eliminates regression to the mean [19, 20, 21].

Control required systolic pressure below 140 mmHg and diastolic pressure below 90 mmHg. Participants were classified as controlled at both visits (CC), controlled then uncontrolled (CU), uncontrolled then controlled (UC), or uncontrolled at both visits (UU). These labels describe two observed visits; they do not establish continuous control between visits. The primary contrast was UC versus UU among initially uncontrolled participants. CU versus CC among initially controlled participants and a model incorporating all four groups were secondary analyses.

Stroke ascertainment and observation time

The outcome was first reported physician-diagnosed stroke after the landmark. We used wave-specific questionnaire items rather than assuming identically named variables across waves. In particular, the 2020 stroke item differed from earlier waves. Verification of a previous response, correction of prior disease history, and a newly affirmative response were distinguished. Later recalled diagnosis years before or within the landmark calendar year were conservatively excluded, as were age-derived diagnosis intervals indicating prior disease or overlapping the landmark. Explicit 2015 corrections to an earlier negative history were also excluded. These procedures improve temporal classification but rely on recall and may themselves introduce selection.

For an event, its time was represented by the interval from the last negative interview to the first affirmative report. Interview months were represented by mid-month; recalled diagnosis years were not assigned an artificial exact event date. An initial affirmative report was retained if later denied, with exclusion of such participants in sensitivity analysis. Four otherwise unrecorded strokes were identified through the 2020 exit proxy module; a sensitivity analysis omitted these proxy events.

Participants without a recorded stroke were right-censored at their last stroke-negative interview. Death without an observed stroke did not imply stroke-free survival to the date of death. A secondary stroke-or-death endpoint used the interval between the last jointly event-free observation and the earliest known endpoint upper bound. Death reporting dates were distinguished from actual death dates. An actual month was used only when explicitly recorded in the solar calendar and chronologically plausible; lunar or unavailable dates used the survey death-report bound. Consequently, neither the stroke model nor the composite analysis estimates a competing-risk cumulative incidence of stroke. Cause-specific and subdistribution hazards target different quantities [22, 23, 24]. The current last-negative-interview censoring scheme is not relabelled as a fully ascertained cause-specific competing-risk analysis. Although interval-censored competing-risk models exist [25], they require additional assumptions about observation and event type; ordinary Fine–Gray fitting after substituting death-report dates for exact death dates would not resolve unknown pre-death stroke status. Death certificates or registry-linked causes were not available as validated endpoint records for this analysis.

Covariates and statistical analysis

Baseline covariates were age, sex, education, rural hukou, body mass index, current smoking, ever drinking alcohol, contemporaneously reported diabetes and broad heart disease, western antihypertensive medication, and continuous systolic and diastolic pressure. Rural hukou refers to registration status and should not be interpreted as current residence. Missing responses were not generally coded as absence of disease or treatment. Follow-up pressure was part of the exposure definition and was not additionally entered as a covariate; follow-up treatment was not adjusted for as if it preceded the exposure.

The primary analysis used an interval-censored Weibull accelerated failure-time model fitted with R survival. Under the Weibull parametrization, coefficients were transformed to proportional hazard ratios as exp(− beta/scale) [26]. Community-clustered covariance was propagated through this transformation using the delta method, including uncertainty in the scale parameter. Complete cases formed the primary clinical model. Confidence intervals were two-sided 95% intervals. The primary contrast was specified in a local analysis plan before association modelling; feasibility and unadjusted event counts had already been examined. This was not a preregistered study. Secondary comparisons were not corrected for multiplicity and are interpreted as exploratory.

Prespecified sensitivity analyses included a midpoint-time Cox model, an interval-censored lognormal model, alternative blood pressure thresholds and averaging rules, exclusion of later stroke denials and exit proxy events, restriction to baseline western medication users, relaxed historical exclusions, baseline biomarker weighting, a 2015 landmark, and multiple imputation. The latter used 20 chained-equation datasets with predictive mean matching and Rubin pooling of log hazard ratios. Its outcome auxiliaries used event status and approximate observation time; the imputation was not a full joint model of interval censoring or community clustering. Chained-equation imputation requires appropriate conditional models and missingness assumptions [27, 28]; compatibility with the substantive survival model is an additional concern [29]. The direction of any resulting bias cannot be assumed to be conservative. Baseline biomarker weights do not turn this selected longitudinal sample into a nationally representative population estimate.

A post hoc piecewise-exponential proportional hazards analysis allowed baseline hazards to differ across 0–2, 2–5, and more than 5 years. A separate post hoc analysis weighted for having any usable post-landmark stroke observation, with stabilized weights truncated at the 1st and 99th percentiles. This addresses initial observation selection under measured predictors, not time-varying informative dropout. Weighting requires measured predictors, positivity and correctly specified observation probabilities [30, 31]. It cannot recover a 2013 control state for participants who died before that visit, and we do not claim transport to every baseline hypertensive participant. These analyses were added after model diagnostics and are explicitly distinguished from the local plan.

Further post-revision analyses fitted a semiparametric interval-censored proportional hazards model with icenReg, avoiding a Weibull baseline shape, and calculated a percentile interval from 300 community-level bootstrap samples [32, 33]. Community resampling retained all individuals in each selected community; the number of clusters, rather than the size of the smallest cluster alone, is relevant to the uncertainty calculation. Robust Cox inference provides related methodological background [34], but does not by itself validate the interval model. Unadjusted Turnbull estimates were used descriptively [35]. Events with intervals wider than four years were treated as censored at the last negative interview in an additional analysis, rather than excluding only their event records without changing observation time.

Illustrative death-related underreporting scenarios retained the original initially uncontrolled complete-case sample. Among known deaths without reported stroke, we randomly assigned hypothetical

unreported strokes within the interval from last negative report to the death upper bound: 10% or 20% in both groups, 20% in UC only, or 20% in UU only. Each nonzero scenario used 100 draws. These probabilities are uncalibrated assumptions. Scenario summaries describe changes in point estimates, not confidence intervals, estimated correction factors, or competing-risk incidence.

Use of AI assistance

OpenAI Codex was used to assist with code development, analysis checks, figure preparation, and English-language drafting and revision. The authors are responsible for the accuracy and integrity of the analysis and manuscript.

Results

Participant characteristics

The complete-case sample included 2902 participants from 406 communities, with a mean age of 61.0 years (standard deviation, 9.2); 1609 (55.4%) were women. There were 488 CC, 288 CU, 823 UC, and 1303 UU participants. Their respective stroke event counts were 33, 33, 54, and 136 (Table 1). These counts correspond to observed report proportions and should not be read as fixed-duration cumulative risks because observation and censoring differed between participants.

Table 1
Participant characteristics by blood pressure control transition

Characteristic	CC	CU	UC	UU	SMD
Participants	488	288	823	1303	—
Reported stroke events	33	33	54	136	—
Age, years	59.3 (8.2)	60.7 (8.7)	60.4 (9.2)	62.0 (9.6)	-0.167
Body mass index, kg/m²	25.0 (4.1)	24.9 (4.3)	24.1 (3.6)	24.4 (4.1)	-0.076
Baseline systolic BP, mmHg	122.4 (10.7)	127.9 (9.4)	150.6 (13.4)	157.6 (15.7)	-0.478
Baseline diastolic BP, mmHg	72.5 (9.0)	74.5 (9.0)	85.7 (10.3)	87.2 (11.3)	-0.145
2013 systolic BP, mmHg	123.3 (10.7)	152.3 (13.7)	126.2 (9.6)	157.2 (15.4)	-2.417
2013 diastolic BP, mmHg	72.4 (8.8)	84.9 (10.7)	73.7 (8.3)	86.6 (11.5)	-1.286
Male	208 (42.6%)	117 (40.6%)	392 (47.6%)	576 (44.2%)	0.069
Rural hukou	375 (76.8%)	214 (74.3%)	677 (82.3%)	1077 (82.7%)	-0.010
Current smoker	119 (24.4%)	65 (22.6%)	264 (32.1%)	396 (30.4%)	0.036
Ever drank alcohol	196 (40.2%)	105 (36.5%)	331 (40.2%)	489 (37.5%)	0.055
Baseline diabetes	68 (13.9%)	38 (13.2%)	48 (5.8%)	82 (6.3%)	-0.019
Baseline heart disease	109 (22.3%)	65 (22.6%)	118 (14.3%)	186 (14.3%)	0.002
Baseline western antihypertensive medication	324 (66.4%)	217 (75.3%)	221 (26.9%)	464 (35.6%)	-0.190
Education: below primary	224 (45.9%)	145 (50.3%)	408 (49.6%)	669 (51.3%)	-0.035
Education: primary	108 (22.1%)	57 (19.8%)	190 (23.1%)	311 (23.9%)	-0.018
Education: secondary plus	156 (32.0%)	86 (29.9%)	225 (27.3%)	323 (24.8%)	0.058
SBP change, mmHg	0.9 (13.0)	24.4 (15.5)	-24.4 (15.6)	-0.4 (18.0)	-1.425

Characteristic	CC	CU	UC	UU	SMD
DBP change, mmHg	-0.2 (9.6)	10.4 (9.9)	-12.0 (10.2)	-0.7 (10.0)	-1.118
Observation endpoint, years	6.4 (1.3)	6.2 (1.5)	6.4 (1.4)	6.2 (1.6)	0.122
2013 western medication, yes	302 (61.9%)	201 (69.8%)	293 (35.6%)	563 (43.2%)	-0.156
2013 western medication, unknown	77 (15.8%)	25 (8.7%)	440 (53.5%)	515 (39.5%)	0.282

Values are mean (standard deviation) or count (percentage). C and U denote measured controlled and uncontrolled blood pressure in 2011 and 2013. SMD compares UC with UU. Changes are 2013 minus 2011. The observation endpoint is not reverse-Kaplan–Meier median follow-up. All listed baseline covariates are complete in this selected sample; missingness before selection appears in Additional file 1. Unknown 2013 medication status is shown explicitly.

Among the 2982 participants with usable stroke follow-up, 262 first reports were recorded, including four exit proxy reports. Forty had a positive report followed by a negative response. The 80 participants excluded from the clinical complete-case analysis had missingness in one or more covariates, most often body mass index, diabetes, or baseline medication. The larger history-cleaned master sample contained 464 known deaths: 61 in CC, 32 in CU, 110 in UC, and 261 in UU. Of these, 444 had no recorded stroke and 20 had a recorded stroke. Only 24 participants in this master sample had an exit-module death year and month; 21 were labelled solar calendar and three lunar calendar. This limited timing information underscores the uncertainty about stroke before death.

Among the 256 complete-case events, interval width had a median of 2.92 years (interquartile range, 2.00–3.00; range, 1.75–5.08), and 11 intervals exceeded four years. Community sizes had a median of seven participants (interquartile range, 3–10; range, 1–26), with 34 singleton communities. Of 6326 baseline hypertensive participants, the 2902 eventually included were younger on average than the 3424 excluded participants (61.0 versus 62.8 years). This comparison combines several eligibility and observation mechanisms and is descriptive.

Mean systolic change was − 24.4 mmHg in UC, compared with + 24.4 mmHg in CU; the groups are defined partly by these changes, so their separation is not evidence of intervention efficacy. Follow-up western medication status was unknown in 440 of 823 UC participants and 515 of 1303 UU participants. Medication initiation, dose intensification, and adherence therefore cannot be reliably inferred from pressure transitions. Full group summaries, missingness, and wave observation counts are in Additional file 1.

Primary and secondary associations

Among 2126 initially uncontrolled participants with 190 events, the adjusted HR comparing UC with UU was 0.683 (95% CI, 0.492–0.948; P = 0.023). The four-group model gave a similar UC-versus-UU association (HR, 0.662; 95% CI, 0.478–0.915), although this model uses a different population and shared covariate effects. Among 776 initially controlled participants with 66 events, CU versus CC was associated with a higher modelled hazard (HR, 1.980; 95% CI, 1.170–3.349). The latter estimate is secondary and has limited event support for the number of adjusted parameters (Table 2; Fig. 2).

Table 2
Main and secondary adjusted associations

Population and contrast	Endpoint	N	Events	HR (95% CI)	P
Initially uncontrolled; UC vs UU	Reported stroke	2126	190	0.683 (0.492–0.948)	0.023
Initially controlled; CU vs CC	Reported stroke	776	66	1.980 (1.170–3.349)	0.011
Four-group sample; UC vs UU	Reported stroke	2902	256	0.662 (0.478–0.915)	0.013
Initially uncontrolled; UC vs UU	Stroke or death	2224	534	0.748 (0.609–0.918)	0.005

Interval-censored Weibull models adjusted for baseline age, sex, education, rural hukou, body mass index, smoking, alcohol use, diabetes, heart disease, western antihypertensive medication and continuous systolic and diastolic pressure. Confidence intervals account for community clustering and scale uncertainty. The initially uncontrolled stroke comparison is primary; others are exploratory. The composite does not estimate stroke cumulative incidence with competing mortality. Full four-group coefficients appear in Additional file 2.

For the stroke-or-death endpoint, the initially uncontrolled complete-case sample comprised 2224 participants and 534 events. The UC-versus-UU HR was approximately 0.75 (95% CI, 0.61–0.92). This association concerns a composite endpoint with broader observation intervals, not stroke incidence after formally accounting for competing mortality.

Sensitivity analyses and model uncertainty

Multiple imputation increased the initially uncontrolled sample to 2192 participants with 195 events and yielded an HR of 0.702 (95% CI, 0.506–0.973). The midpoint Cox estimate was approximately 0.69 (95% CI, 0.49–0.95). Its proportional-hazards diagnostic did not detect a group departure within the primary population (P = 0.209), although the group diagnostic in the four-group sample did (P = 0.036). These diagnostics depend on the midpoint approximation and do not establish the validity of the interval model.

Changing the threshold to 130/80 mmHg produced an HR of approximately 0.76 (95% CI, 0.53–1.09). Averaging all valid readings produced an HR of 0.727 (95% CI, 0.526–1.004). Restriction to baseline western medication users gave an HR of approximately 0.70 (95% CI, 0.42–1.16). The direction was similar, but these intervals included the null. The 2015 landmark analysis gave an HR of approximately 0.68 (95% CI, 0.47–0.98); it is an overlapping-cohort sensitivity analysis, not external validation (Table 3; Fig. 3).

Table 3
Sensitivity analyses among initially uncontrolled participants

Analysis	N	Events	HR (95% CI)	P
Primary Weibull	2126	190	0.683 (0.492–0.948)	0.023
Midpoint Cox approximation	2126	190	0.686 (0.494–0.952)	0.024
Biomarker-weighted	2126	190	0.688 (0.475–0.995)	0.047
Exclude later stroke denials	2101	165	0.685 (0.485–0.969)	0.032
Exclude exit proxy events	2126	186	0.685 (0.491–0.954)	0.025
Baseline medication users	685	79	0.699 (0.422–1.157)	0.164
130/80 mmHg threshold	2491	216	0.757 (0.527–1.088)	0.133
Mean of all valid readings	2091	192	0.727 (0.526–1.004)	0.053
Retain landmark-year diagnoses	2143	207	0.722 (0.526–0.992)	0.045
Report-only outcome definition	2198	262	0.744 (0.560–0.989)	0.042
2015 landmark	1983	147	0.680 (0.471–0.981)	0.039
Multiple imputation	2192	195	0.702 (0.506–0.973)	0.034
Piecewise PH (post hoc)	2126	190	0.685 (0.492–0.952)	0.024
Observation weighting (post hoc)	2126	190	0.685 (0.493–0.950)	0.023

Contrast is UC versus UU. Sample definitions differ across rows. The all-reading upper confidence bound exceeds one. The lognormal time ratio is presented separately in Fig. 4 and Results. The 2015 cohort overlaps the primary cohort. Further revision-stage models and death-underreporting scenarios appear in Additional file 1. None of these estimates is a causal treatment effect.

The lognormal interval model yielded a time ratio of 1.22 (95% CI, 0.97–1.53), where a ratio above one represents a longer modelled time to stroke. Its Akaike information criterion was lower than that of the Weibull model in the primary population (approximately 1542 versus 1551). Time ratios and hazard ratios are different estimands; this finding nevertheless shows that the strength of the evidence depends on distributional assumptions. Post hoc piecewise-exponential modelling yielded an HR of

approximately 0.68 (95% CI, 0.49–0.95). Agreement among selected proportional-hazards models does not remove the uncertainty identified by the lognormal analysis.

The semiparametric interval-censored PH estimate was 0.686 (community-bootstrap 95% interval, 0.463–0.959), with all 300 bootstrap fits completing. The distinct hazard-ratio and time-ratio scales are shown separately in Fig. 4. Restriction to seated and fully compliant measurements at both visits gave an HR of 0.702 (95% CI, 0.499–0.988; 1985 participants, 181 events). Censoring wide-interval events at their last negative interview gave an HR of 0.710 (95% CI, 0.506–0.997). Continuous decreases of 10 mmHg systolic pressure and 5 mmHg diastolic pressure, modelled separately with baseline pressure adjustment, gave HRs of 0.922 (95% CI, 0.857–0.992) and 0.901 (95% CI, 0.843–0.963), respectively. These are post-revision associations and do not constitute measurement-error correction.

The original primary sample contained 71 UC and 175 UU participants eligible for the hypothetical pre-death stroke scenarios. Equal assumed underreporting fractions yielded median HRs near 0.69. Assuming an unreported stroke in 20% of eligible UC deaths but none of the eligible UU deaths shifted the median HR to 0.857; reversing that assumption shifted it to 0.553. These point-estimate shifts demonstrate dependence on differential underascertainment. They do not establish the true direction or magnitude of bias, and their simulation quantiles are not sampling confidence intervals.

Discussion

In this landmark analysis, transition from uncontrolled to controlled measured blood pressure was associated with a lower subsequent reported stroke hazard in the primary Weibull model. The association was similar after multiple imputation and several alternative analyses. However, some exposure definitions yielded intervals including the null, and a lognormal model with better relative fit gave an uncertain time ratio. The evidence therefore supports a cautious observational association rather than a model-independent conclusion.

The finding is directionally compatible with randomized evidence on blood pressure lowering [1], but the exposure in this study is not an assigned intervention. Treatment changes, adherence, access to care, diet, intercurrent disease, and unmeasured vascular risk may influence both control transitions and stroke. Adjustment for baseline pressure helps account for initial severity but cannot fully resolve confounding or regression to the mean. In particular, measurement error at a high initial reading can create an apparent improvement at the second visit without a comparable change in underlying usual pressure.

The contribution is narrower than a general claim that blood pressure predicts cardiovascular outcomes. Cumulative pressure and hypertension care categories have already been investigated in CHARLS [11, 12]. Here, separating the two-visit exposure window from follow-up and reconstructing historical stroke responses focuses the analysis on a clinically interpretable transition. The outcome audit also demonstrates why later affirmative disease reports should not automatically be classified as incident

events. The analysis makes an incremental contribution by clarifying exposure timing and outcome ascertainment; it does not establish a new biological mechanism.

The higher hazard associated with loss of control among initially controlled participants may be clinically relevant, but this secondary analysis contained only 66 events and a relatively rich adjustment set. It should not be interpreted as confirming an asymmetric effect of gaining versus losing control. The different populations and the absence of a formal causal interaction analysis preclude that conclusion. Similarly, the composite result does not establish that stroke reduction explains a mortality association.

Several limitations are consequential. Stroke was self-reported or proxy-reported without event adjudication, and stroke subtype was not resolved. First report and first occurrence need not coincide. Historical exclusions rely on recalled dates, and conservative exclusion of a diagnosis within the landmark year can remove true post-landmark events. Positive responses later denied introduce additional classification uncertainty, although exclusion of these records gave a similar direction. Interval-censored models acknowledge incomplete event timing but still require assumptions about observation and censoring that may fail when attendance depends on illness.

Requiring two pressure assessments selects survivors and returning participants. Later death and loss to follow-up may further select the observable sample. Exact death timing was usually unavailable, and the available weighting analysis cannot correct all informative censoring. Measured control at two visits cannot establish sustained pressure control, and medication classes and adherence were not adequately characterized. Visit-to-visit variability has been studied using richer repeated-measurement designs [36]; two observations here do not provide a reliable measure of long-term variability. Within-visit replicates capture only part of measurement error and cannot identify all between-visit error or white-coat effects. Without a suitable validation study, formal calibration would require unverified error assumptions [21]. The observational sample was not externally validated in another cohort; the other locally available databases were not incorporated into this analysis. Finally, the plan was locally recorded after feasibility inspection, several analyses were exploratory, and the lognormal result limits any assertion of comprehensive robustness.

For community hypertension surveillance, the relevant question is whether repeat measurements of control status provide prognostic information beyond a single assessment. The present association supports further investigation in cohorts with adjudicated stroke and more complete treatment and mortality records; it does not establish that a monitoring programme prevents stroke. These data alone do not identify a treatment strategy, justify a 140/90 versus 130/80 target, or support an individual absolute risk reduction. In target-trial terms, attaining a biomarker threshold is not itself a sufficiently specified assignable treatment strategy [37]. We therefore retain an observational landmark design rather than relabel it as an emulation of treatment initiation or intensification. Independent replication should retain a prespecified landmark and outcome definition rather than select a favourable exposure threshold after seeing results.

Conclusions

In the primary Weibull model, attaining measured control between two visits was associated with a lower modelled hazard of subsequently reported stroke among initially uncontrolled participants. A semiparametric model gave a similar direction, but the evidence remained sensitive to distributional assumptions and exposure definitions, with several analyses unable to exclude no association. Uncertain stroke status before death, measurement error, and selection preclude a causal treatment interpretation.

Abbreviations

ACCORD, Action to Control Cardiovascular Risk in Diabetes; AFT, accelerated failure time; AIC, Akaike information criterion; BMI, body mass index; BP, blood pressure; CC, controlled at both visits; CHARLS, China Health and Retirement Longitudinal Study; CI, confidence interval; CU, controlled to uncontrolled; DBP, diastolic blood pressure; HR, hazard ratio; PH, proportional hazards; PROGRESS, Perindopril Protection Against Recurrent Stroke Study; SBP, systolic blood pressure; SMD, standardized mean difference; SPRINT, Systolic Blood Pressure Intervention Trial; STEP, Strategy of Blood Pressure Intervention in the Elderly Hypertensive Patients; STROBE, Strengthening the Reporting of Observational Studies in Epidemiology; UC, uncontrolled to controlled; UU, uncontrolled at both visits.

Declarations

Ethics approval and consent to participate

The China Health and Retirement Longitudinal Study (CHARLS) was approved by the Institutional Review Board of Peking University (approval numbers IRB00001052-11015 for the household survey, including anthropometric assessments, and IRB00001052-11014 for biomarker collection). The original CHARLS study was conducted in accordance with the Declaration of Helsinki and relevant institutional ethical guidelines. Written informed consent was obtained from all participants in the original surveys. The present study was a secondary analysis of de-identified CHARLS data, involved no new participant recruitment or intervention, and did not access direct personal identifiers. Therefore, no additional institutional ethics approval or additional informed consent was required for the present secondary analysis.

Consent for publication

Not applicable.

Competing interests

The authors declare that they have no competing interests.

Additional files

Additional file 1 (.docx). Supplementary methods, reporting checklist and results. Contains the questionnaire crosswalk, measurement and mortality definitions, STROBE checklist, supplementary tables and Figure S1.

Additional file 2 (.zip). Analysis code and aggregate results. Contains Python and R scripts, the locally recorded analysis plan, machine-readable summary tables and reproducibility instructions. No individual-level source or derived records are included.

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Author Contribution

BBZ and YY contributed to conceptualisation. BBZ, SAUS and ZL contributed to methodology. BXZ and YQ contributed to data curation and validation. LZ and MG contributed to visualisation and validation. BBZ drafted the manuscript. SAUS, ZL, BL and KZ contributed to manuscript review and editing. BL, KZ and YY provided supervision. YY was responsible for project administration and funding acquisition. Author initials: BBZ, Binbin Zhang; BXZ, Baixin Zheng; YQ, Yongjia Qiang; SAUS, Shafiu A Umar Shinge; ZL, Zhuxuan Liu; LZ, Lu Zhang; MG, Minnan Gao; BL, Bin Li; KZ, Kuan Zeng; YY, Yanqi Yang. Binbin Zhang and Shafiu A Umar Shinge contributed equally to this work and share first authorship. All authors have read and approved the final manuscript and its submission to BMC Public Health.

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Availability of data and materials

The CHARLS datasets analysed in this study are available through the official CHARLS data-access service [38], subject to the provider's access conditions. The analysis used the 2011, 2013, 2015, 2018, and 2020 survey releases and Harmonized CHARLS version D. Supplementary methods, results, and the STROBE checklist are included in Additional file 1. Aggregate analysis outputs and Python and R scripts

are included in Additional file 2. Individual-level source and derived data are not redistributed and should be obtained or reconstructed under the data provider's access conditions.

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Figures

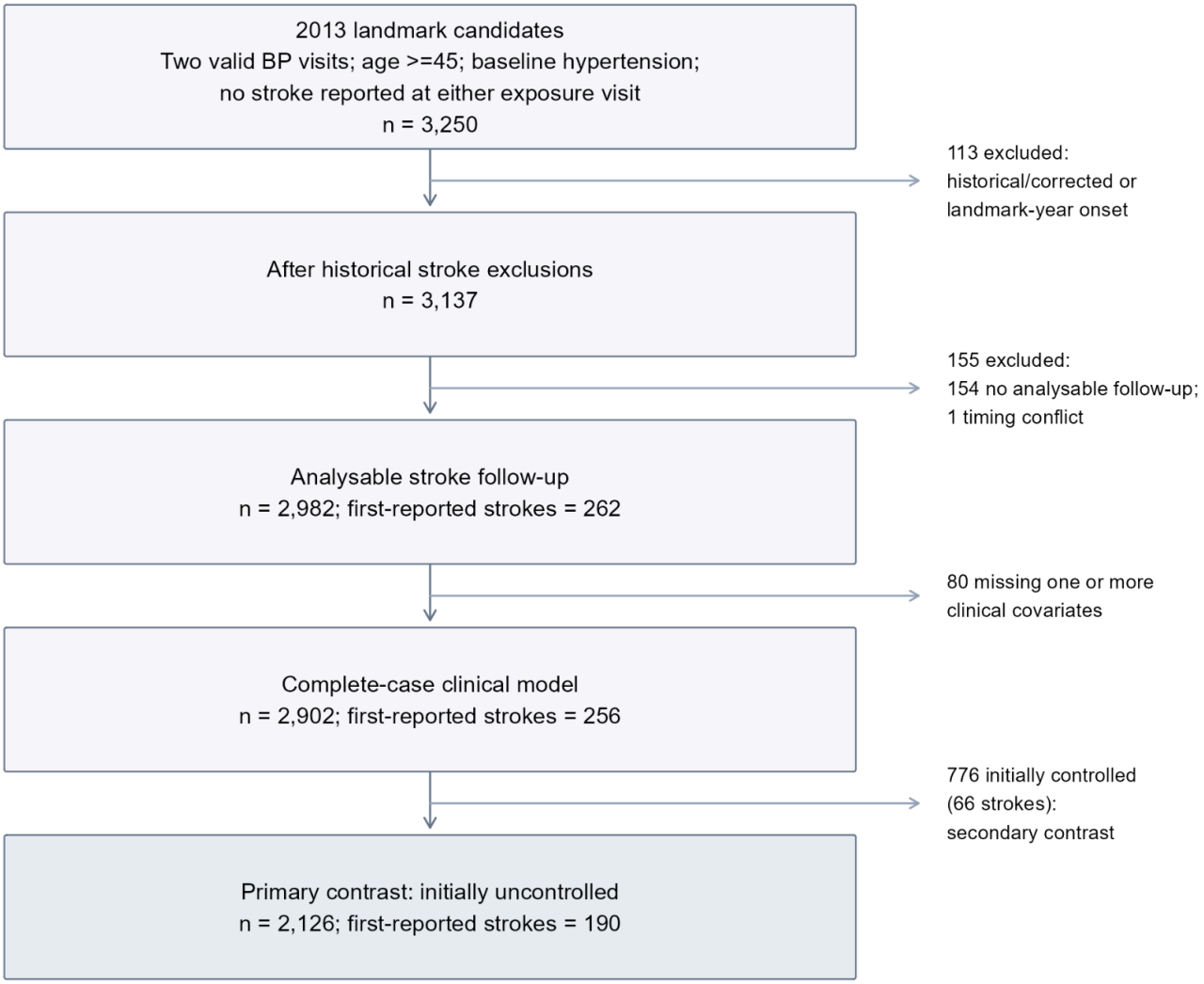


Figure 1

Construction of the 2013 landmark sample. Historical exclusions use later recalled diagnosis timing and explicit corrections to previous negative stroke history. Complete-case exclusions refer to clinical covariates, after determining usable outcome observation. The primary comparison is restricted to initially uncontrolled participants. BP control required systolic pressure below 140 mmHg and diastolic pressure below 90 mmHg. Counts are participants, not person-years.

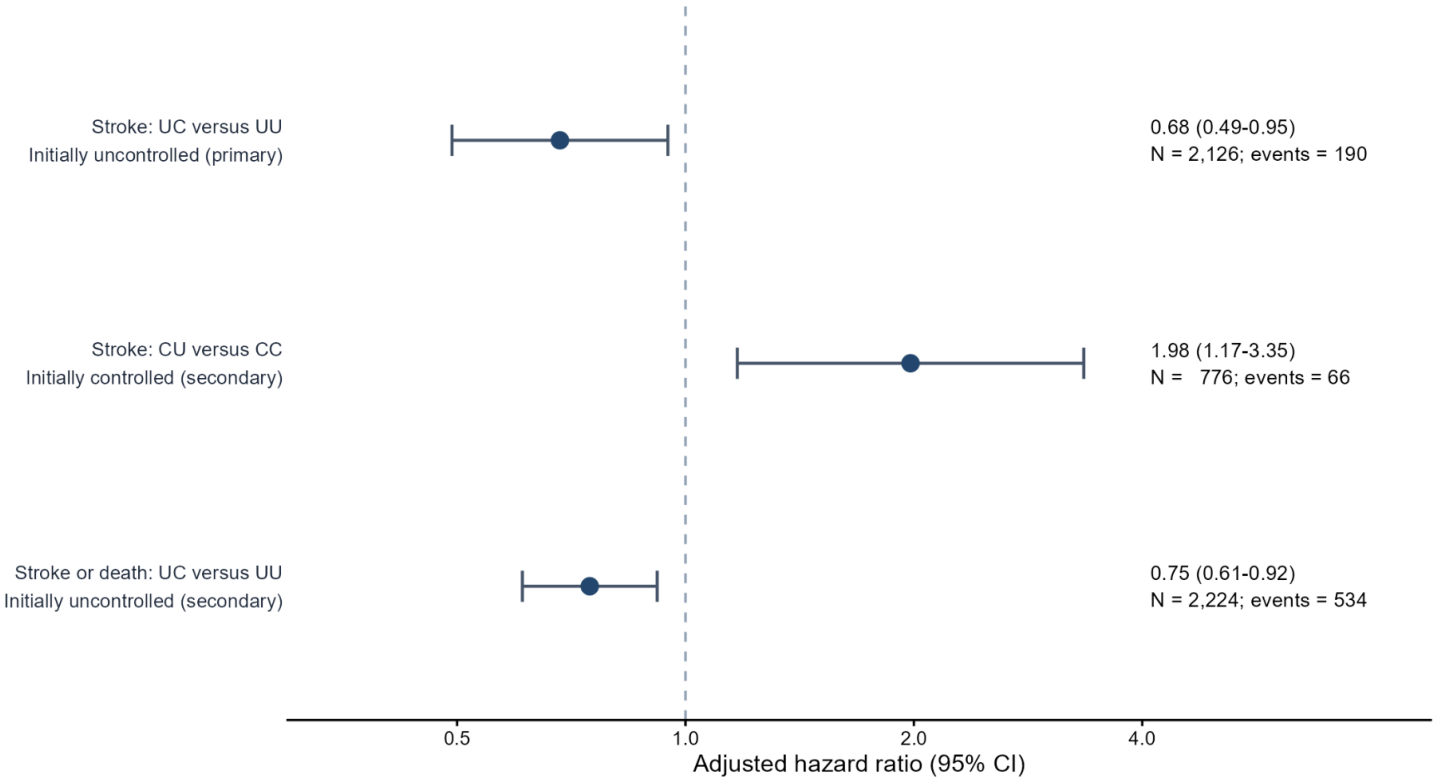


Figure 2

Main and secondary associations. Points show adjusted hazard ratios and horizontal lines show 95% confidence intervals from interval-censored Weibull models, with community-clustered covariance and scale uncertainty included. Each row uses its indicated population and endpoint. The primary stroke contrast includes 2126 participants and 190 events; the initially controlled secondary contrast includes 776 and 66; the initially uncontrolled composite contrast includes 2224 and 534. Secondary comparisons are exploratory and not multiplicity-adjusted. The vertical line denotes HR = 1.

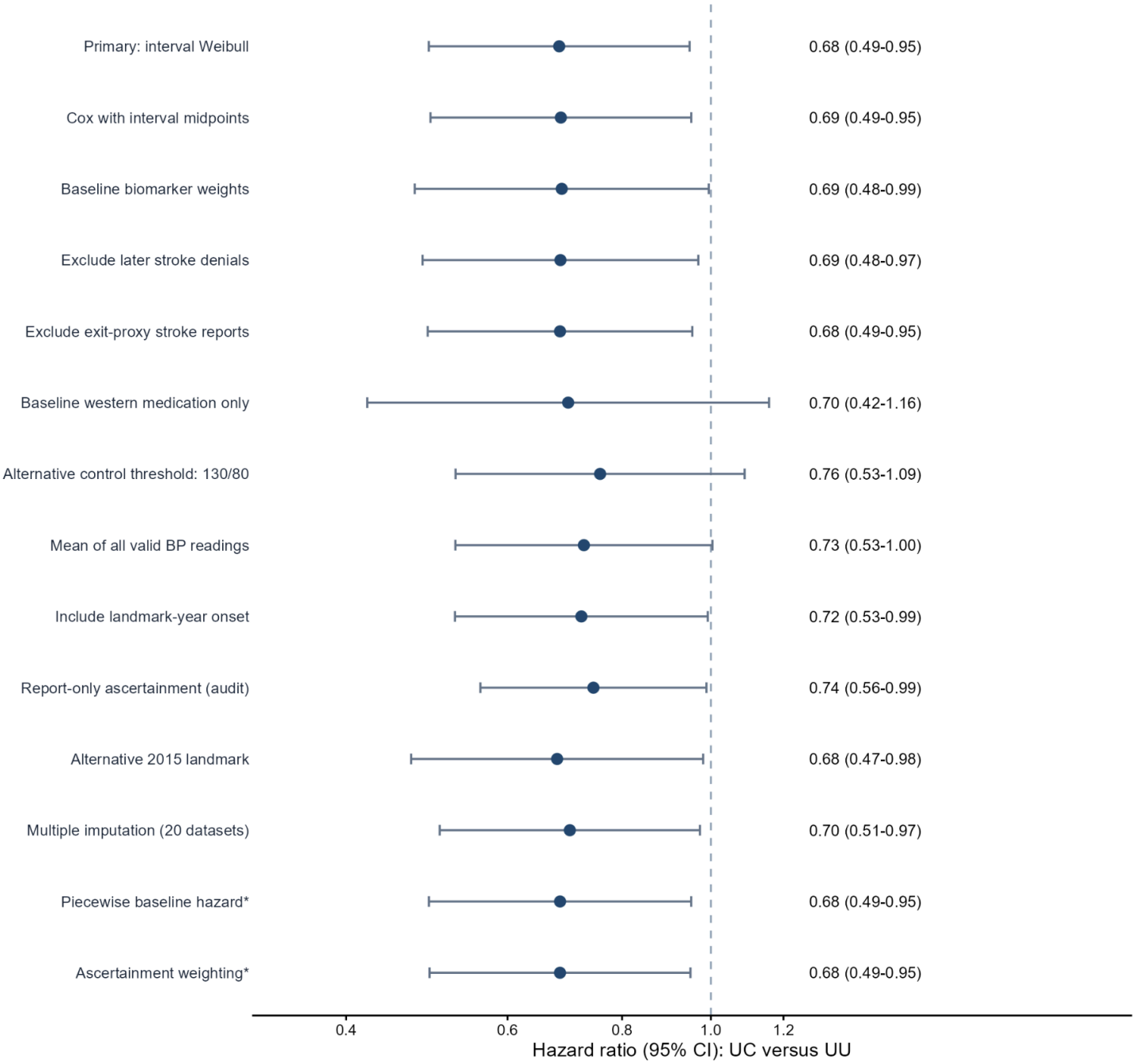


Figure 3

Sensitivity analyses for UC versus UU among initially uncontrolled participants. Points and lines indicate hazard ratios and 95% confidence intervals. Exposure definitions and sample restrictions can change the population. Post hoc analyses are explicitly marked. The all-valid-reading upper confidence limit is 1.004 before rounding and includes the null. The lognormal result is reported in the text as a time ratio and is not placed on the hazard-ratio axis. No plotted association represents a causal treatment effect. Asterisks identify diagnostic additions after initial modelling.

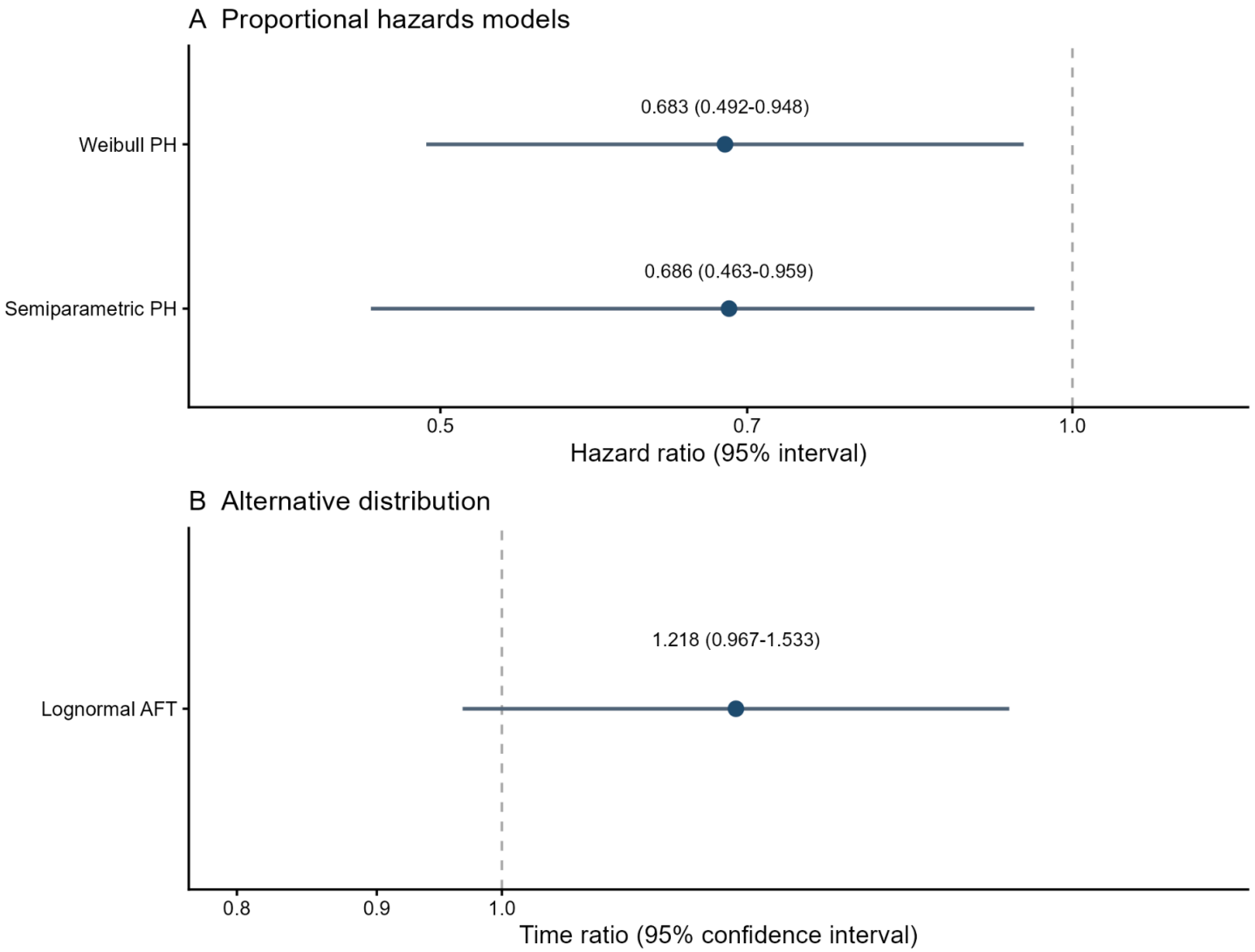


Figure 4

Distributional sensitivity on separate effect scales. Panel A shows the primary Weibull and semiparametric interval-censored proportional hazards estimates. The Weibull interval uses community-clustered Wald uncertainty; the semiparametric interval is a percentile interval from 300 community bootstrap samples. Panel B shows the lognormal accelerated failure-time estimate as a time ratio. The null is one on both axes, but the estimands differ. This presentation is not a test of the difference between models.

Supplementary Files

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