

Association of Pre-Stroke Dietary Inflammatory Potential with Admission Neurological Deficit and 90-Day Functional Outcome after Acute Ischemic Stroke

Jinan Fang

Second Affiliated Hospital of Harbin Medical University

Peijia Liu

Second Affiliated Hospital of Harbin Medical University

Fan Wang

Harbin Medical University

Maoqing Wang

Harbin Medical University

Ying Cao

Second Affiliated Hospital of Harbin Medical University

Zhaojun Liu

Second Affiliated Hospital of Harbin Medical University

Xuesong Sun

Second Affiliated Hospital of Harbin Medical University

Siyi Chen

Second Affiliated Hospital of Harbin Medical University

Zhaobo Zhang

Second Affiliated Hospital of Harbin Medical University

Xingbang Zou

Second Affiliated Hospital of Harbin Medical University

Yin Huang

China Pharmaceutical University

Peifang Liu

liupeifang@hrbmu.edu.cn

Second Affiliated Hospital of Harbin Medical University

Keywords: Acute ischemic stroke, Dietary inflammatory index, Functional outcome, Modified Rankin Scale, Neurological deficit, Inflammation, Diet

Posted Date: October 11th, 2026

DOI: <https://doi.org/10.21203/rs.3.rs-11009762/v1>

License: © ⓘ This work is licensed under a Creative Commons Attribution 4.0 International License.
[Read Full License](#)

Additional Declarations: No competing interests reported.

Abstract

Background

The Dietary Inflammatory Index (DII) summarizes the inflammatory potential of habitual diet, but evidence linking pre-stroke dietary inflammatory potential to neurological injury and recovery after acute ischemic stroke (AIS) remains limited.

Methods

We retrospectively analyzed a consecutively enrolled single-center observational cohort of patients with AIS recruited from June 1, 2024 through May 2025. Habitual diet during the year before stroke was assessed using a semi-quantitative food-frequency questionnaire, and a literature-derived DII was calculated from 30 available non-energy dietary parameters. The primary outcome was unfavorable functional outcome approximately 90 days after hospital admission (modified Rankin Scale [mRS] 2–6 vs 0–1; permitted assessment window 90 ± 7 days after hospital admission); the secondary outcome was admission National Institutes of Health Stroke Scale (NIHSS) > 3 . Multivariable logistic regression, restricted cubic splines, and sensitivity analyses were performed.

Results

Among 303 patients, 106 (35.0%) had a 90-day unfavorable functional outcome and 112 (37.0%) had admission NIHSS > 3 . In 299 complete cases, each 1-SD higher DII was associated with greater odds of the 90-day outcome after adjustment for demographic factors, vascular risk factors, and reperfusion therapy (OR 1.70, 95% CI 1.29–2.25; $P < 0.001$), and the association remained after additional adjustment for admission NIHSS (OR 1.59, 95% CI 1.20–2.11; $P = 0.001$). The adjusted OR for admission NIHSS > 3 was 1.34 (95% CI 1.04–1.74; $P = 0.026$) per 1-SD higher DII. The alternative mRS 3–6 endpoint and restriction to first-ever stroke supported the primary association. The 90-day spline model supported an overall association without establishing statistically significant nonlinearity.

Conclusions

Higher pre-stroke dietary inflammatory potential was associated with greater neurological deficit at admission and worse 90-day functional outcome after AIS. These observational findings require confirmation in larger, multicenter cohorts with standardized dietary assessment, prospectively retained energy and anthropometric data, and broader representation of stroke severity.

Background

Acute ischemic stroke (AIS) remains a major cause of death and long-term disability worldwide. Despite advances in reperfusion therapy and secondary prevention, substantial variation in neurological injury and functional recovery remains after stroke, and a meaningful proportion of stroke burden cannot be explained by established vascular risk factors alone [1, 2]. Identifying modifiable exposures that precede the acute event may therefore improve understanding of interindividual differences in post-stroke recovery and may inform future prevention-oriented research.

Inflammation is central to ischemic brain injury. Cerebral ischemia triggers danger-associated signaling, microglial activation, leukocyte recruitment, endothelial dysfunction, blood-brain barrier disruption, and systemic immune responses that can influence infarct evolution and neurological recovery [3, 4]. Diet is one of several long-term environmental exposures capable of modulating systemic inflammatory and oxidative pathways [5–7]. Rather than focusing on single nutrients, the Dietary Inflammatory Index (DII) was developed to quantify the overall inflammatory potential of an individual diet on the basis of published associations between dietary components and inflammatory biomarkers [5]. Higher scores indicate a more pro-inflammatory dietary pattern, whereas lower scores indicate a more anti-inflammatory pattern.

Most epidemiological studies of DII and cerebrovascular disease have focused on the occurrence of stroke or related vascular phenotypes. Analyses from population-based datasets have reported associations between higher DII and prevalent or incident stroke, while a study in patients with ischemic stroke linked higher DII to greater carotid plaque vulnerability [8–10]. More recently, a large UK Biobank analysis linked higher energy-adjusted DII to incident ischemic stroke in women and identified related metabolic signatures [23]. These findings support the biological relevance of dietary inflammatory potential to cerebrovascular disease, but they primarily address pre-event risk rather than neurological injury at presentation or functional recovery after AIS. Evidence addressing these post-event clinical outcomes remains comparatively limited.

Dietary inflammation is also linked to metabolic dysfunction and insulin resistance [11, 12]. The triglyceride-glucose index (TyG), calculated from fasting triglyceride and glucose concentrations, is a pragmatic surrogate marker of insulin resistance and has been studied in relation to stroke severity and outcome, with heterogeneous findings across cohorts and meta-analyses [13–16]. Because inflammatory and metabolic pathways may interact, it is reasonable to explore whether TyG modifies associations between DII and stroke outcomes; however, such analyses should be considered exploratory rather than predictive.

We therefore investigated the association of pre-stroke habitual dietary inflammatory potential with 90-day functional outcome in patients with AIS, with admission neurological deficit severity as a secondary outcome. We additionally examined dose-response patterns and robustness under alternative outcome definitions and a first-ever stroke restriction. Correlations with inflammatory biomarkers, DII-only discrimination, and possible effect modification by TyG were treated as exploratory analyses rather than as predictive or mechanistic tests.

Methods

Study design and setting

This study was a retrospective analysis of a consecutively enrolled single-center observational cohort conducted at the Second Affiliated Hospital of Harbin Medical University. Consecutive hospitalized patients with AIS were screened for eligibility and recruited from June 1, 2024 through May 2025. The manuscript was prepared in accordance with the Strengthening the Reporting of Observational Studies in Epidemiology (STROBE) recommendations for observational studies and the STROBE-nut extension for nutritional epidemiology [21, 22].

The study protocol, "Association of the Dietary Inflammatory Index and Triglyceride-Glucose Index with the Severity and Prognosis of Acute Ischemic Stroke," was approved by the Ethics Committee of the Second Affiliated Hospital of Harbin Medical University on May 29, 2024 (approval no. YJSKY2024-292; approved study period May 2024 to May 2026). Recruitment for the present cohort began on June 1, 2024, after ethics approval, and all participants included in this analysis were recruited through May 2025.

Written informed consent for participation was obtained from all included participants in accordance with the approved procedures at the time of enrollment. The present study retrospectively analyzed data from this observational cohort and associated clinical records.

Study participants

Patients were eligible if they met the diagnostic criteria for AIS in the Chinese Guidelines for Diagnosis and Treatment of Acute Ischemic Stroke 2018 [18], were 18–80 years of age, were admitted within 72 hours of symptom onset, were able to participate in the dietary assessment, received standardized acute-phase and secondary-prevention treatment, and provided informed consent. Patients with intellectual disability, psychiatric illness, cognitive impairment, disturbance of consciousness, aphasia, or another condition that prevented completion of the questionnaire were not eligible for the dietary assessment.

Exclusion criteria were malignant tumor; severe cardiac, pulmonary, hepatic, renal, or gastrointestinal disease; acute or chronic infection; immune dysfunction; use of antibiotics or corticosteroids; implausible estimated energy intake (< 800 or > 4200 kcal/day for men and < 500 or > 3500 kcal/day for women); progressive ischemic stroke; pre-stroke mRS score ≥ 2 . Participant selection is summarized in Fig. 1.

Of 442 patients assessed for eligibility, 71 were excluded at the initial eligibility-screening stage. Among the remaining 371 potentially eligible patients, 68 were excluded after application of the additional exclusion criteria, leaving 303 patients who completed the dietary assessment. No participant was lost

to follow-up or had missing 90-day outcome data; therefore, all 303 participants had 90-day mRS data available and formed the descriptive cohort. Figure 1 presents the aggregate numbers at each screening stage together with the applicable exclusion criteria.

Dietary assessment

Habitual dietary intake during the year preceding stroke onset was assessed within 48 hours of admission using a semi-quantitative food-frequency questionnaire (SQFFQ) adapted to Chinese dietary habits [17]. The questionnaire covered staple grains and tubers, legumes and soy products, vegetables, mushrooms and sea vegetables, fruits, meat and poultry, fish and shellfish, eggs, dairy products, nuts, and related food groups. Two investigators who had completed standardized online training conducted face-to-face interviews. A food atlas and standard portion information were used to support recall of usual intake. Interviews were conducted with the patient or, when necessary, with a cohabiting informant who was familiar with the patient's habitual diet. The identity of the respondent as either the patient or an informant was not systematically retained in the analytical dataset; therefore, self-reported and informant-assisted dietary assessments could not be distinguished in subsequent analyses. The questionnaire also asked about calcium, iron, vitamin, and other supplement use during the preceding 3 months; because supplement dose information was not collected in sufficient detail, supplement-derived nutrient amounts were not incorporated into the DII calculation.

For each food item, participants were asked to report usual portion size and consumption frequency. When a participant could provide an approximate specific frequency, that reported frequency was used. When only a predefined frequency range could be recalled, the midpoint of the category was used: 4–6 times/week was converted to 5 times/week, 1–3 times/week to 2 times/week, and 1–3 times/month to 2 times/month; daily intake was converted to once per day and never or rarely consumed foods were assigned a frequency of zero. Daily food intake was estimated by combining portion size with the corresponding daily-equivalent consumption frequency.

Dietary component intakes were estimated primarily using the China Food Composition Tables, Standard Edition, 6th edition [19]. When required values were unavailable from the Chinese source, supplementary food-composition values in the formal calculation template were obtained from United States Department of Agriculture food-composition data. The exact USDA database release/access year was not retained in the analytical records.

Calculation of the Dietary Inflammatory Index

The DII was calculated using the literature-derived approach described by Shivappa et al. [5]. For each available dietary parameter, the participant's estimated daily intake was standardized using the corresponding global reference mean and standard deviation to generate a Z score. The Z score was converted to a centered percentile score and multiplied by the published inflammatory effect score for

that parameter. The global reference values and inflammatory effect scores were taken from the DII development framework described by Shivappa et al. [5]. Parameter-specific scores were then summed to obtain the overall DII; higher values represent a more pro-inflammatory dietary pattern.

For the final analytical dataset, the energy parameter was excluded from the DII component set, and DII was calculated from 30 non-energy dietary parameters. The parameters, their units, global reference values, and inflammatory effect scores are presented in Supplementary Table S1.

The analysis used a literature-derived DII calculated from 30 available non-energy dietary parameters rather than the energy-adjusted DII (E-DII). Estimated total energy intake was used only during dietary-data screening to apply the prespecified sex-specific plausibility criteria. Energy itself was not included as a DII component, and participant-level total-energy values were not retained in the final analytical dataset; therefore, total energy could not be included as an additional covariate and E-DII could not be calculated.

Clinical and laboratory data

Demographic and clinical variables included sex, age, systolic and diastolic blood pressure, reported smoking history, reported alcohol drinking, hypertension, diabetes mellitus, hyperlipidemia, coronary heart disease, atrial fibrillation, previous stroke, intravenous thrombolysis, mechanical thrombectomy, and admission NIHSS score. Reperfusion therapy was defined as receipt of intravenous thrombolysis and/or mechanical thrombectomy during the index hospitalization. The questionnaire also requested height, weight, and educational attainment; however, a substantial proportion of participants could not provide sufficiently reliable responses for these items, so BMI and education were not included in the analytical dataset or regression models.

Because the study used binary history variables obtained mainly from patient/family report, the smoking and alcohol variables should be interpreted as reported smoking history and reported alcohol drinking rather than rigorously quantified current exposure or cumulative dose.

Peripheral blood-cell measurements were obtained on the day of admission. Fasting triglyceride (TG) and fasting blood glucose (FBG) were measured from peripheral venous blood after at least 8 hours of fasting, generally on the second hospital day. Acute clinical treatment was not delayed for research sampling; consequently, fasting TG and FBG were obtained after acute treatment in some patients. All laboratory measurements were obtained as part of routine clinical care through the hospital clinical laboratory.

Specific analyzer models, manufacturers, reagent details, and instrument-level quality-control information were not retained in the research dataset and are therefore not reported.

TG and FBG values measured in mmol/L were converted to mg/dL using conversion factors of 88.57 and 18.0, respectively, before calculating TyG as $\ln[\text{TG (mg/dL)} \times \text{FBG (mg/dL)} / 2]$. Inflammatory composite

indices were calculated as follows: neutrophil-to-lymphocyte ratio (NLR) = neutrophil count/lymphocyte count; platelet-to-lymphocyte ratio (PLR) = platelet count/lymphocyte count; systemic immune-inflammation index (SII) = platelet count x neutrophil count/lymphocyte count; and systemic inflammation response index (SIRI) = monocyte count x neutrophil count/lymphocyte count.

Outcome assessment

Initial neurological deficit was assessed immediately at admission by neurologists with more than 3 years of clinical experience using the NIHSS. The prespecified secondary binary outcome was moderate-to-severe neurological deficit, defined as NIHSS > 3 versus ≤ 3. This cutoff was selected to distinguish very mild neurological deficits (NIHSS ≤ 3) from greater deficit burden while retaining a clinically interpretable binary secondary endpoint. NIHSS was also analyzed descriptively as a continuous/ordinal score across DII quartiles.

The primary clinical outcome was functional status assessed by telephone approximately 90 days after hospital admission, with a permitted assessment window of 90 ± 7 days after hospital admission, using the modified Rankin Scale (mRS). Follow-up was performed by neurologists with more than 3 years of clinical experience. During the telephone interview, the assessor asked about the patient's current daily activities, degree of independence, and living status and assigned the corresponding mRS score. The primary binary definition of unfavorable functional outcome was mRS 2–6 versus 0–1, selected to contrast excellent functional recovery (mRS 0–1) with any residual disability. Death was coded as mRS 6. The full mRS distribution from 0 to 6 was retained for ordinal analysis. To assess robustness to a more dependency-oriented definition, mRS 3–6 versus 0–2 was evaluated as an alternative binary endpoint.

No formal assessor-blinding procedure was used for the 90-day telephone follow-up; follow-up assessors were aware of participants' clinical information. Accordingly, this study is not described as single-blind.

Covariates

Covariates were selected on clinical and epidemiological grounds rather than solely by univariable statistical significance. For the 90-day outcome, Model 3 adjusted for age, sex, reported smoking history, reported alcohol drinking, hypertension, diabetes mellitus, atrial fibrillation, coronary heart disease, previous stroke, and reperfusion therapy. Model 4 additionally adjusted for admission NIHSS to examine whether the association persisted after conditioning on neurological severity at presentation. Because admission NIHSS may partly lie on the pathway linking pre-stroke exposures to later functional status, Model 3 and Model 4 were interpreted together rather than treating NIHSS adjustment as simply additional confounding control. For the neurological-deficit outcome, the corresponding adjusted model included the Model 3 covariates but not admission NIHSS because NIHSS defined the outcome. NLR

was not included in the primary adjustment sets because systemic inflammation could plausibly lie on the pathway linking dietary inflammatory potential to stroke outcomes.

Statistical analysis

Analyses were organized according to the study objectives. The primary inferential analysis evaluated the association between continuous DII, expressed per 1-SD increase, and the 90-day unfavorable functional outcome. The secondary inferential analysis evaluated the association between DII and moderate-to-severe neurological deficit at admission. Analyses based on DII quartiles, restricted cubic splines, alternative definitions of functional outcome, the ordinal mRS scale, and restriction to patients with first-ever stroke were considered supportive or sensitivity analyses. Analyses of DII-TyG interaction, correlations with inflammatory biomarkers, and DII-only discrimination were considered exploratory.

Continuous variables were summarized as median (interquartile range) and categorical variables as number (percentage). Baseline characteristics across DII quartiles were compared using the Kruskal-Wallis rank-sum test for continuous variables and Fisher's exact test for categorical variables. DII was modeled continuously per 1-unit increase, per 1-standard-deviation (SD) increase (1 SD = 1.762 DII units), and by quartiles. Trend tests across quartiles used the median DII of each quartile entered as a continuous variable in regression models.

For descriptive Fig. 2 analyses, NIHSS distributions across DII quartiles were compared using the Kruskal-Wallis test and trend was assessed using Spearman rank correlation with ordinal quartile number. The full mRS 0–6 distribution was compared using Fisher's exact test with 50000 Monte Carlo replicates, with Spearman correlation used for trend. The unadjusted trend in the proportion with mRS 2–6 was assessed using binary logistic regression with DII quartile entered as an ordinal variable.

Associations of DII with the two binary outcomes were evaluated using logistic regression. Model 1 was unadjusted. Model 2 adjusted for age, sex, reported smoking history, and reported alcohol drinking. Model 3 additionally adjusted for hypertension, diabetes mellitus, atrial fibrillation, coronary heart disease, previous stroke, and reperfusion therapy. For the 90-day outcome only, Model 4 additionally adjusted for admission NIHSS as a severity-adjusted analysis. Odds ratios (ORs) and 95% confidence intervals (CIs) were reported. Variance-inflation factors in the most extensively adjusted covariate set were < 1.5, providing no evidence of problematic multicollinearity.

Restricted cubic spline (RCS) logistic regression was used to explore dose-response shape, with three knots at the 10th, 50th, and 90th percentiles of DII and the median DII (0.506) as the reference. P-overall was obtained by comparing the covariate-only model with the spline model, and P-nonlinear by comparing the linear DII model with the spline model using likelihood-ratio tests. For the 90-day outcome, the spline model used the Model 4 adjustment set, including admission NIHSS; for neurological deficit, it used the Model 3 covariate set without NIHSS.

Sensitivity analyses for the 90-day outcome included an alternative binary endpoint (mRS 3–6 vs 0–2), a proportional-odds cumulative logit model across the full mRS 0–6 range, and restriction to patients with first-ever stroke. For the first-ever stroke analysis, previous stroke was omitted from the covariate set. The common OR from the ordinal model represents the odds of being in a worse mRS category per 1-SD increase in DII. The proportional-odds assumption for DII was assessed using a likelihood-ratio test comparing the proportional-odds model with a model allowing a nominal (threshold-specific) DII effect. Because evidence of non-proportionality was detected, additional exploratory cumulative binary logistic models were fitted at successive mRS thresholds (mRS 1–6 vs 0, 2–6 vs 0–1, 3–6 vs 0–2, 4–6 vs 0–3, 5–6 vs 0–4, and 6 vs 0–5), using the same covariate adjustment set as the primary 90-day outcome model. Threshold-specific estimates based on very sparse events or accompanied by separation warnings were considered unstable and were not interpreted as reliable effect estimates.

A fully non-proportional global diagnostic model was also explored; because that model did not converge, the global likelihood-ratio statistic was not used for inference. Interpretation of the ordinal analysis therefore focused on the DII-specific proportional-odds diagnostic and the exploratory threshold-specific cumulative models.

Exploratory effect-modification analyses were performed in complementary forms. For the continuous interaction analysis reported in Table 3, DII and TyG were standardized to z scores and their product term was entered into the outcome-specific logistic model; P for interaction was obtained from a likelihood-ratio test comparing nested models with and without the DII x TyG interaction term. Separately, DII and TyG were dichotomized at their medians (DII 0.506; TyG 8.808), with values equal to the median classified as low, to describe joint exposure groups (Table S4). Joint-category analyses were fitted with binary logistic regression, and global joint-group and multiplicative-interaction P values were obtained from likelihood-ratio tests. For the 90-day outcome, models adjusted for age, sex, reported smoking history, reported alcohol drinking, hypertension, diabetes mellitus, atrial fibrillation, coronary heart disease, previous stroke, reperfusion therapy, and admission NIHSS; for the neurological-deficit outcome, the same covariates were used except NIHSS. Additive interaction was evaluated using modified Poisson regression with a log link and HC0 robust standard errors. RERI and AP were calculated from adjusted relative risks, with 95% CIs obtained by the delta method; the multiplicative-interaction P value in the modified Poisson model was based on the robust Wald test. Correlations between DII and inflammatory biomarkers were assessed using two-sided Spearman correlations with pairwise complete observations, Benjamini-Hochberg false-discovery-rate (FDR) adjustment, and percentile-bootstrap 95% CIs based on 2000 resamples. Exploratory receiver-operating-characteristic (ROC) analyses evaluated DII alone, without adjustment for clinical covariates; AUC 95% CIs were estimated using the DeLong method.

The same outcome-specific clinical covariate sets were used for the DII-TyG interaction analyses. BMI, education, and participant-level total energy intake were not included because sufficiently reliable or retained values were unavailable for the final analytical dataset.

Analyses used complete cases for the covariates required by each model. Four patients had missing information on previous stroke and were excluded from multivariable models requiring that covariate, leaving 299 participants. Other biomarker analyses used available observations as indicated. No total-energy-adjusted model was performed because participant-level energy values were not retained in the final analytic dataset. Two-sided $P < 0.05$ was considered statistically significant. All statistical analyses were performed in R version 4.4.0 (R Foundation for Statistical Computing, Vienna, Austria).

Use of generative artificial intelligence in manuscript preparation

During manuscript preparation, the authors used ChatGPT (OpenAI) for language and structural editing. All scientific content, statistical analyses, interpretations, and final wording were critically reviewed and approved by the authors.

Results

Participant selection and analytical cohorts

A total of 442 hospitalized patients were assessed for eligibility. After 71 first-stage exclusions, 371 patients remained potentially eligible; a further 68 patients were excluded after application of the additional exclusion criteria. The final descriptive cohort included 303 patients who completed the SQFFQ. No participant was lost to follow-up or had missing 90-day outcome data, and 90-day mRS data were available for all 303 participants (Fig. 1).

Figure 1 reports aggregate exclusion counts for the two screening stages and lists the corresponding exclusion criteria. Individual reason-specific counts were not available for reliable reconstruction and are therefore not presented separately.

Four participants had missing previous-stroke information and were excluded from models requiring that covariate; therefore, the complete-case analytical cohort included 299 participants. In this cohort, 105 had a 90-day unfavorable functional outcome and 110 had moderate-to-severe neurological deficit at admission.

[Figure 1. Participant selection and analytical cohorts]

Baseline characteristics across DII quartiles

In the full cohort, median age was 62 years (IQR 56–69), 217 participants (71.6%) were male, median admission NIHSS was 3 (IQR 1–5), and median DII was 0.51 (IQR – 0.89 to 1.73). The DII distribution had a mean of 0.384, SD of 1.762, median of 0.506, and range from – 4.078 to 4.360. Quartile groups included 76, 76, 76, and 75 participants, respectively.

Several baseline characteristics differed across DII quartiles (Table 1). The proportion of women increased from 13.2% in Q1 to 42.7% in Q4 ($P < 0.001$), whereas reported alcohol drinking decreased from 46.1% to 22.7% ($P = 0.014$). Admission NIHSS scores differed across quartiles ($P = 0.020$). High-sensitivity C-reactive protein, triglycerides, serum uric acid, and TyG also differed across quartiles, although these unadjusted comparisons were descriptive and were not used to select covariates.

Table 1
Baseline characteristics according to DII quartiles

Variable	Overall	Q1	Q2	Q3	Q4	P value
Dietary inflammatory index	0.51 (-0.89, 1.73)	-1.69 (-2.36, -1.38)	-0.19 (-0.60, 0.21)	1.20 (0.84, 1.43)	2.36 (1.92, 3.07)	< 0.001
Female	86 (28.4%)	10 (13.2%)	16 (21.1%)	28 (36.8%)	32 (42.7%)	< 0.001
Male	217 (71.6%)	66 (86.8%)	60 (78.9%)	48 (63.2%)	43 (57.3%)	
Age, years	62.00 (56.00, 69.00)	60.00 (56.00, 66.00)	63.50 (54.50, 69.50)	62.50 (56.50, 69.50)	61.00 (56.00, 68.00)	0.454
Systolic blood pressure, mmHg	149.00 (136.00, 168.00)	147.50 (135.50, 170.00)	152.50 (139.50, 169.50)	149.50 (134.50, 165.00)	150.00 (134.00, 167.00)	0.478
Diastolic blood pressure, mmHg	90.00 (80.00, 97.00)	89.00 (79.00, 99.50)	91.00 (81.00, 99.00)	88.00 (79.50, 93.50)	90.00 (82.00, 97.00)	0.210
Reported smoking history	117 (38.6%)	33 (43.4%)	32 (42.1%)	25 (32.9%)	27 (36.0%)	0.500
Reported alcohol drinking	100 (33.0%)	35 (46.1%)	27 (35.5%)	21 (27.6%)	17 (22.7%)	0.014
Hypertension	158 (52.1%)	46 (60.5%)	41 (53.9%)	36 (47.4%)	35 (46.7%)	0.286
Diabetes mellitus	76 (25.1%)	20 (26.3%)	15 (19.7%)	23 (30.3%)	18 (24.0%)	0.511
Coronary heart disease	20 (6.6%)	3 (3.9%)	7 (9.2%)	6 (7.9%)	4 (5.3%)	0.577
Atrial fibrillation	11 (3.6%)	1 (1.3%)	5 (6.6%)	2 (2.6%)	3 (4.0%)	0.346
Hyperlipidemia	11 (3.6%)	2 (2.6%)	2 (2.6%)	2 (2.6%)	5 (6.7%)	0.536
Previous stroke	84 (28.1%)	24 (32.4%)	20 (26.3%)	23 (30.7%)	17 (23.0%)	0.566
Missing	4	2	0	1	1	
Intravenous thrombolysis	192 (63.4%)	47 (61.8%)	46 (60.5%)	47 (61.8%)	52 (69.3%)	0.660
Mechanical thrombectomy	3 (1.0%)	0 (0.0%)	1 (1.3%)	1 (1.3%)	1 (1.3%)	0.904
Admission NIHSS score	3.00 (1.00, 5.00)	2.00 (1.00, 4.00)	2.00 (1.00, 5.00)	3.00 (1.00, 6.00)	3.00 (2.00, 6.00)	0.020

Variable	Overall	Q1	Q2	Q3	Q4	P value
Platelet count	222.00 (190.00, 268.00)	217.50 (186.50, 251.50)	211.50 (175.00, 258.00)	225.00 (192.00, 271.50)	234.00 (202.00, 273.00)	0.150
Neutrophils, %	73.00 (64.90, 82.10)	68.95 (62.95, 79.70)	73.20 (65.65, 81.95)	74.25 (64.35, 83.65)	75.00 (68.10, 83.70)	0.050
High-sensitivity C-reactive protein	2.34 (1.19, 5.18)	1.77 (0.69, 3.76)	2.39 (1.19, 7.54)	3.93 (1.88, 7.57)	2.08 (1.13, 3.81)	0.007
Missing	58	14	18	11	15	
Neutrophil-to-lymphocyte ratio	3.37 (2.24, 5.75)	3.12 (2.14, 5.03)	3.43 (2.36, 5.00)	3.33 (2.13, 5.75)	3.70 (2.59, 6.58)	0.244
Platelet-to-lymphocyte ratio	137.00 (102.78, 187.68)	120.97 (96.42, 186.14)	130.38 (102.87, 181.21)	136.22 (98.67, 202.58)	143.27 (116.83, 183.33)	0.148
Systemic immune-inflammation index	768.65 (493.70, 1346.56)	681.01 (464.17, 1175.04)	719.80 (491.56, 1212.52)	785.67 (460.32, 1316.86)	848.19 (627.71, 1629.38)	0.073
Systemic inflammation response index	0.96 (0.53, 1.79)	0.85 (0.45, 1.57)	1.07 (0.67, 1.85)	0.97 (0.52, 1.67)	0.95 (0.57, 2.01)	0.466
Fasting blood glucose	5.59 (4.95, 6.87)	5.78 (5.17, 6.71)	5.27 (4.85, 5.92)	5.57 (4.94, 6.92)	5.63 (4.95, 7.12)	0.157
Triglycerides	1.42 (1.06, 2.15)	1.66 (1.24, 2.65)	1.39 (0.99, 2.22)	1.44 (1.07, 2.00)	1.24 (1.00, 1.85)	0.018
Homocysteine	12.15 (9.73, 16.03)	12.21 (9.09, 14.96)	12.38 (10.16, 15.99)	12.86 (9.57, 17.06)	12.00 (9.55, 16.57)	0.874
Missing	9	4	3	0	2	
Serum uric acid	319.60 (258.50, 393.10)	335.05 (278.70, 401.55)	321.95 (279.15, 395.95)	325.55 (279.10, 407.95)	275.00 (220.50, 335.80)	< 0.001
Triglyceride-glucose index	8.81 (8.42, 9.41)	8.96 (8.60, 9.53)	8.61 (8.31, 9.18)	8.86 (8.41, 9.39)	8.61 (8.38, 9.18)	0.022
Albumin	43.40 (40.70, 45.60)	43.95 (41.25, 45.50)	42.00 (40.70, 44.80)	43.30 (39.20, 45.60)	43.20 (41.10, 46.40)	0.488
Missing	66	12	25	17	12	

Values are median (Q1, Q3) or n (%). Percentages for variables with missing observations were calculated using available observations as the denominator. P values are from Kruskal-Wallis rank-sum tests for continuous variables and Fisher exact tests for categorical variables. Abbreviations: DII, Dietary Inflammatory Index; NIHSS, National Institutes of Health Stroke Scale.

DII and 90-day functional outcome

At 90 days, 197 patients (65.0%) had mRS 0–1 and 106 (35.0%) had an unfavorable outcome (mRS 2–6). The full mRS distribution was: mRS 0, n = 112; mRS 1, n = 85; mRS 2, n = 49; mRS 3, n = 39; mRS 4, n = 11; mRS 5, n = 4; and mRS 6, n = 3. Median DII was 0.29 (IQR – 1.13 to 1.44) in patients with mRS 0–1 and 1.04 (IQR – 0.51 to 2.23) in those with mRS 2–6 ($P < 0.001$; Table S2).

In unadjusted descriptive analyses, the proportion with mRS 2–6 increased across DII quartiles from 26.3% (20/76) in Q1 to 31.6% (24/76) in Q2, 32.9% (25/76) in Q3, and 49.3% (37/75) in Q4 (P for trend = 0.0046; Fig. 2C). The overall distribution of mRS 0–6 across quartiles did not reach conventional statistical significance in the Monte Carlo Fisher test ($P = 0.0526$), although the ordinal trend was significant (Spearman $P = 0.0043$; Fig. 2B).

In multivariable logistic regression, the association with 90-day unfavorable functional outcome was present before and after additional adjustment for admission NIHSS. In Model 3, each 1-SD increase in DII was associated with higher odds of unfavorable outcome (OR 1.70, 95% CI 1.29–2.25; $P < 0.001$). After additional adjustment for admission NIHSS in Model 4, the corresponding OR was 1.59 (95% CI 1.20–2.11; $P = 0.001$), equivalent to an OR of 1.30 (95% CI 1.11–1.53; $P = 0.001$) per 1-unit increase in DII. In Model 4, compared with Q1, adjusted ORs were 1.12 (95% CI 0.53–2.39) for Q2, 1.18 (95% CI 0.55–2.54) for Q3, and 2.48 (95% CI 1.16–5.34; $P = 0.020$) for Q4, with P for trend = 0.027 (Table 2).

Dose-response association with 90-day outcome

The three-knot RCS model for the 90-day outcome, adjusted using the Model 4 covariate set, showed an overall association between DII and unfavorable functional outcome (P -overall < 0.001 ; Fig. 3A). The curve suggested that odds rose more steeply at higher DII values, but statistical evidence for a nonlinear component did not meet the conventional 0.05 threshold (P -nonlinear = 0.060). Accordingly, the spline analysis was interpreted as supporting an overall association without establishing a statistically significant nonlinear threshold or clinical cut point.

DII and initial neurological deficit

At admission, 191 patients (63.0%) had NIHSS ≤ 3 and 112 (37.0%) had NIHSS > 3 . Median DII was 0.25 (IQR – 1.11 to 1.50) in the mild-deficit group and 0.88 (IQR – 0.49 to 1.90) in the moderate-to-severe group ($P = 0.012$; Table S3). NIHSS distributions differed across DII quartiles (Kruskal-Wallis $P = 0.0199$), with a significant ordinal trend (Spearman $P = 0.0018$; Fig. 2A).

[Figure 2. Stroke severity and 90-day outcomes across DII quartiles]

After full adjustment, the OR for moderate-to-severe neurological deficit was 1.18 (95% CI 1.02–1.37; $P = 0.026$) per 1-unit increase in DII and 1.34 (95% CI 1.04–1.74; $P = 0.026$) per 1-SD increase. Compared with Q1, the adjusted OR for Q4 was 2.54 (95% CI 1.23–5.27; $P = 0.012$), with P for trend = 0.023 (Table 2). In contrast to the prespecified linear model, the higher-degree-of-freedom RCS model did not show a statistically significant overall spline association (P -overall = 0.076) and provided no evidence of nonlinearity (P -nonlinear = 0.828; Fig. 3B).

Table 2

Associations between DII and clinical outcomes Part A. Association between DII and 90-day unfavorable functional outcome

DII expression	Model 1	Model 2	Model 3	Model 4
DII, per 1-unit increase	1.30 (1.13–1.51) $P < 0.001$	1.34 (1.15–1.56) $P < 0.001$	1.35 (1.16–1.58) $P < 0.001$	1.30 (1.11–1.53) $P = 0.001$
DII, per 1-SD increase	1.60 (1.24–2.06) $P < 0.001$	1.68 (1.28–2.19) $P < 0.001$	1.70 (1.29–2.25) $P < 0.001$	1.59 (1.20–2.11) $P = 0.001$
Q1	1.00 (reference)	1.00 (reference)	1.00 (reference)	1.00 (reference)
Q2	1.29 (0.64–2.61) $P = 0.475$	1.29 (0.63–2.63) $P = 0.480$	1.30 (0.62–2.71) $P = 0.485$	1.12 (0.53–2.39) $P = 0.768$
Q3	1.37 (0.68–2.76) $P = 0.375$	1.46 (0.71–2.99) $P = 0.301$	1.43 (0.68–2.99) $P = 0.347$	1.18 (0.55–2.54) $P = 0.670$
Q4	2.73 (1.38–5.39) $P = 0.004$	3.02 (1.48–6.16) $P = 0.002$	3.06 (1.46–6.42) $P = 0.003$	2.48 (1.16–5.34) $P = 0.020$
P for trend	0.006	0.003	0.004	0.027

Values are ORs (95% CIs). Model 1 unadjusted. Model 2: age, sex, reported smoking history, reported alcohol drinking. Model 3: Model 2 plus hypertension, diabetes mellitus, atrial fibrillation, coronary heart disease, previous stroke, and reperfusion therapy. Model 4: Model 3 plus admission NIHSS. Model 4 used 299 complete cases and should be interpreted as an association conditional on admission neurological severity rather than as a total-effect estimate. P for trend used quartile-specific median DII entered continuously.

Part B. Association between DII and initial neurological deficit

DII expression	Model 1	Model 2	Model 3
DII, per 1-unit increase	1.16 (1.01–1.33) P = 0.031	1.17 (1.02–1.35) P = 0.028	1.18 (1.02–1.37) P = 0.026
DII, per 1-SD increase	1.30 (1.02–1.65) P = 0.031	1.33 (1.03–1.70) P = 0.028	1.34 (1.04–1.74) P = 0.026
Q1	1.00 (reference)	1.00 (reference)	1.00 (reference)
Q2	1.62 (0.82–3.20) P = 0.169	1.57 (0.79–3.14) P = 0.201	1.55 (0.76–3.16) P = 0.231
Q3	1.36 (0.68–2.72) P = 0.381	1.38 (0.68–2.80) P = 0.379	1.42 (0.69–2.94) P = 0.343
Q4	2.42 (1.23–4.76) P = 0.011	2.54 (1.25–5.16) P = 0.010	2.54 (1.23–5.27) P = 0.012
P for trend	0.024	0.021	0.023

Values are ORs (95% CIs). Model 1 unadjusted. Model 2: age, sex, reported smoking history, reported alcohol drinking. Model 3: Model 2 plus hypertension, diabetes mellitus, atrial fibrillation, coronary heart disease, previous stroke, and reperfusion therapy. Admission NIHSS was not included because NIHSS defined the outcome. Analyses used complete cases.

[Figure 3. Dose-response associations of DII with stroke outcomes]

Sensitivity analyses

The primary association was supported by the alternative binary endpoint and the first-ever stroke restriction (Table 3; Figure S3). For mRS 3–6 versus 0–2, the fully adjusted OR per 1-SD higher DII was 1.93 (95% CI 1.33–2.80; $P < 0.001$). Among 215 patients with first-ever stroke, 77 had the primary unfavorable outcome, and the corresponding adjusted OR was 1.79 (95% CI 1.28–2.49; $P < 0.001$). In the proportional-odds model across mRS 0–6, the common OR for being in a worse mRS category was 1.35 (95% CI 1.08–1.70; $P = 0.010$); however, the proportional-odds assumption was not supported for DII (DII-specific likelihood-ratio test, $P = 0.025$). Exploratory threshold-specific cumulative analyses therefore showed that the association was not uniform across the mRS scale (Table S8): ORs per 1-SD higher DII were 1.15 (95% CI 0.89–1.49; $P = 0.279$) for mRS 1–6 vs 0, 1.59 (95% CI 1.20–2.11; $P = 0.001$) for mRS 2–6 vs 0–1, 1.93 (95% CI 1.33–2.80; $P < 0.001$) for mRS 3–6 vs 0–2, and 1.59 (95% CI 0.82–3.07; $P = 0.167$) for mRS 4–6 vs 0–3. Estimates for mRS 5–6 vs 0–4 (7 events) and death vs survival (3 deaths) were unstable because of sparse events and possible separation and were not interpreted.

Exploratory TyG and interaction analyses

In continuous interaction analyses, there was no statistically significant evidence that TyG modified the DII association with either outcome (Table 3). For the 90-day unfavorable outcome, the interaction OR per 1-SD DII x 1-SD TyG was 0.78 (95% CI 0.59–1.02; P for interaction = 0.072); for moderate-to-severe neurological deficit, the interaction OR was 0.97 (95% CI 0.75–1.25; P for interaction = 0.802). Median-based joint analyses likewise did not provide consistent evidence of multiplicative interaction (Table S4).

Table 3
Sensitivity and exploratory interaction analyses of DII and clinical outcomes

Analysis	Outcome	N/events	Effect estimate (95% CI)	P value	P for interaction
Part A. Sensitivity analyses					
Primary definition	90-day mRS 2–6 vs 0–1	299/105	1.59 (1.20–2.11)	0.001	
Alternative definition	90-day mRS 3–6 vs 0–2	299/56	1.93 (1.33–2.80)	< 0.001	
Ordinal mRS	90-day mRS 0–6	299	Common OR 1.35 (1.08–1.70)	0.010	
First-ever stroke only	90-day mRS 2–6 vs 0–1	215/77	1.79 (1.28–2.49)	< 0.001	
Part B. Exploratory continuous DII x TyG interaction analyses					
DII x TyG interaction	90-day mRS 2–6 vs 0–1	299/105	Interaction OR 0.78 (0.59–1.02)	0.073	0.072
DII x TyG interaction	Admission NIHSS > 3 vs ≤ 3	299/110	Interaction OR 0.97 (0.75–1.25)	0.801	0.802

Part A effect estimates are ORs per 1-SD increase in DII. For the ordinal outcome, the common OR represents the odds of being in a worse mRS category, but the DII-specific proportional-odds assumption was not supported (P = 0.025); the common OR should therefore be interpreted cautiously, with threshold-specific cumulative results provided in Table S8. Primary and alternative binary models adjusted for age, sex, reported smoking history, reported alcohol drinking, hypertension, diabetes mellitus, atrial fibrillation, coronary heart disease, previous stroke, reperfusion therapy, and admission NIHSS; the first-ever stroke model omitted previous stroke. For Part B, DII and TyG were standardized before multiplication; P for interaction was obtained from a likelihood-ratio test comparing nested models with and without the interaction term.

Modified Poisson analyses did not provide evidence of additive interaction (Table S5). For the 90-day unfavorable outcome, RERI was – 0.72 (95% CI -1.66 to 0.22; P = 0.135) and AP was – 0.58 (95% CI -1.32 to 0.16; P = 0.124). For neurological deficit, RERI was 0.27 (95% CI -0.36 to 0.90; P = 0.407) and AP was 0.19 (95% CI -0.26 to 0.65; P = 0.404). These analyses were exploratory and did not support a reproducible synergistic effect between DII and TyG.

Correlations with inflammatory biomarkers and exploratory discrimination

Spearman analyses showed only weak correlations between DII and circulating inflammatory measures (Figure S1). After Benjamini-Hochberg correction, higher DII was positively correlated with neutrophil percentage ($\rho = 0.152$; FDR-adjusted $P = 0.045$), PLR ($\rho = 0.145$; FDR-adjusted $P = 0.045$), and SII ($\rho = 0.152$; FDR-adjusted $P = 0.045$). Correlations with NLR and high-sensitivity C-reactive protein did not remain statistically significant after FDR adjustment. Because these biomarkers were measured during the acute stroke admission, the correlations were interpreted as exploratory rather than evidence of mediation.

DII alone showed limited unadjusted discrimination for the clinical outcomes: the AUC was 0.617 for 90-day unfavorable functional outcome and 0.587 for moderate-to-severe neurological deficit (Figure S2). These supplementary ROC analyses were descriptive only and were not intended to establish a prediction model, validate a clinical threshold, or support clinical decision-making.

Discussion

In this single-center cohort of patients with AIS, a more pro-inflammatory habitual diet before stroke was associated with both worse 90-day functional status and greater neurological deficit at admission. For the 90-day outcome, each 1-SD increase in DII was associated with 70% higher odds after adjustment for demographic factors, vascular risk factors, and reperfusion therapy (Model 3); the association remained after additional conditioning on admission NIHSS (Model 4: 59% higher odds). Because admission NIHSS may be an intermediate between pre-stroke exposures and later disability, the two models address related but not identical questions and should be interpreted together. The secondary NIHSS analysis showed a similar direction of association. The alternative mRS 3–6 endpoint and first-ever stroke restriction supported the primary finding. Ordinal analyses suggested that the association was not uniform across all mRS thresholds, while the spline analysis supported an overall relationship with the 90-day outcome without establishing statistically significant nonlinearity. Exploratory DII-TyG analyses did not demonstrate statistically significant effect modification.

These results extend prior literature by shifting attention from stroke occurrence to post-event clinical outcomes. Existing DII studies have largely evaluated stroke risk or vascular phenotypes [8–10, 23]. Recent AIS cohorts have also shown that nutritional and inflammatory status measured during hospitalization is associated with 90-day functional outcome [24, 25], but such indices reflect acute or post-stroke nutritional-inflammatory state rather than habitual dietary inflammatory potential before the event. By evaluating pre-stroke DII together with admission NIHSS and 90-day mRS, the present study addresses whether the dietary milieu preceding AIS is associated with neurological injury at presentation and subsequent functional status. This distinction may be relevant because post-stroke outcomes depend on the acute lesion and treatment as well as host characteristics that precede the event.

Several biological pathways could plausibly connect habitual dietary inflammatory potential to stroke severity and recovery, although the present observational data cannot establish mechanism. Dietary patterns with greater pro-inflammatory potential have been linked to systemic low-grade inflammation and oxidative stress [6, 7]. Chronic inflammatory signaling can influence endothelial function, atherosclerotic activity, thrombosis, and metabolic homeostasis, while the acute ischemic event itself activates neuroimmune pathways, peripheral leukocyte responses, and blood-brain barrier injury [3, 4]. A pre-existing pro-inflammatory milieu could therefore plausibly modify the host response to cerebral ischemia or the capacity for repair. The weak positive correlations of DII with neutrophil percentage, PLR, and SII in this cohort are directionally compatible with this hypothesis; however, these blood markers were collected after stroke onset and can be strongly affected by the acute event. They should not be interpreted as mediators of the DII-outcome association. Future studies with pre-stroke or serial inflammatory measurements are needed to test mediation explicitly.

The shape of the DII-outcome association also deserves cautious interpretation. The three-knot spline analysis yielded a strong overall association for the 90-day outcome, but the nonlinear component had $P = 0.060$. Although the fitted curve appeared steeper at higher DII values, confidence intervals widened in the tails where fewer observations were available. Therefore, the data do not justify a threshold or clinical cutoff. For initial neurological deficit, the linear and quartile models were statistically significant, whereas the spline model was not significant overall. The additional degrees of freedom and limited sample size likely reduced precision, reinforcing the need to treat the spline analysis as exploratory rather than as a competing primary model.

The ordinal mRS analyses further suggest that the DII-outcome association should not be interpreted as a uniform shift across the entire disability scale. The strongest and most precise threshold-specific associations occurred at cut points distinguishing mRS 0–1 from 2–6 and mRS 0–2 from 3–6. In contrast, the mRS 1–6 vs 0 and mRS 4–6 vs 0–3 estimates were not statistically significant, and the two most severe thresholds were too sparse for reliable estimation. This pattern may indicate that dietary inflammatory potential is more clearly associated with transitions from favorable or mildly impaired states to clinically unfavorable disability than with every increment of the mRS. However, the threshold-specific analyses were exploratory and event counts were small at the severe end of the scale; the findings therefore do not establish the absence of an association with severe disability or mortality and require replication in larger cohorts.

We did not find convincing evidence that TyG modified the association between DII and either outcome. This should not be interpreted as evidence that insulin resistance is irrelevant to stroke. Prior studies and meta-analyses have reported associations of TyG with ischemic stroke occurrence, severity, thrombolysis outcomes, or mortality, but estimates have varied across populations [13–16, 20]. In the present study, TyG was calculated from a single fasting TG and glucose measurement obtained during the acute hospitalization, and some samples were collected after acute treatment. Acute stress hyperglycemia, changes in nutritional intake, fluid therapy, glucose-lowering treatment, and other early clinical interventions can make a single acute-phase TyG value an imperfect proxy for long-term pre-

stroke insulin resistance [16]. The absence of interaction may also reflect limited statistical power. Replication with pre-stroke metabolic data or repeated measurements would provide a stronger test of metabolic effect modification.

The study has several strengths. Dietary exposure was specifically anchored to habitual intake during the year preceding stroke rather than to post-stroke dietary behavior, thereby reducing the possibility that disability-related dietary changes after the index event defined the exposure. The study also evaluated two clinically distinct outcomes: neurological deficit at admission and functional status at 90 days. DII was analyzed continuously, per SD, by quartiles, and with RCS; the 90-day association was examined both before and after additional adjustment for admission NIHSS and was supported by an alternative binary endpoint and a first-ever stroke analysis. Ordinal and threshold-specific analyses were used to characterize potential heterogeneity across the mRS scale rather than to force a single common effect. In addition, DII-only ROC analyses were kept supplementary and were not used to characterize DII as a clinical prediction tool.

Several limitations require emphasis. First, this was a retrospective analysis of a modest-sized, single-center cohort from northeastern China, which limits generalizability to other geographic regions, dietary cultures, populations, and stroke-care settings. In addition, eligibility for the dietary assessment required participants to be able to complete the questionnaire, either directly or with appropriate assistance, and patients with major disturbance of consciousness, cognitive impairment, aphasia, or other conditions preventing questionnaire completion were not eligible. Because these clinical features are associated with stroke severity, severity-related selection bias may have occurred in addition to reduced generalizability. Severe AIS was therefore likely under-represented in the analytical cohort, as reflected by the relatively low median admission NIHSS, the small number of patients undergoing mechanical thrombectomy, and the very small numbers of patients with mRS 5–6 or death at 90 days. The association with admission neurological deficit should consequently be interpreted primarily in patients who were sufficiently stable to undergo dietary assessment. Second, habitual dietary intake during the preceding year was assessed during the acute hospitalization and was therefore subject to recall error. Some assessments were completed with assistance from cohabiting informants, but respondent type was not systematically retained in the analytical dataset. Because the likelihood of requiring informant assistance may itself be related to stroke severity, measurement error in the dietary exposure may not have been entirely nondifferential, and differential exposure misclassification cannot be excluded. Third, estimated total energy intake was used for dietary plausibility screening but participant-level energy values were not retained in the final analytical dataset. Consequently, an energy-adjusted DII (E-DII) could not be calculated and total energy intake could not be included as an additional covariate. Fourth, the DII was calculated from 30 available non-energy dietary parameters rather than all parameters in the original DII framework, which may limit comparability with studies using a broader component set. Fifth, although height, weight, and educational attainment were included in the questionnaire, sufficiently reliable data were not available for inclusion in the final analyses. BMI and educational attainment therefore could not be incorporated into the adjustment models, leaving the possibility of residual confounding related to body size and socioeconomic position. Reported smoking history, alcohol

drinking, and several comorbidities were also based primarily on information provided by patients and/or family members. Additional residual confounding from stroke subtype, lesion location and volume, acute physiological status, and rehabilitation intensity cannot be excluded. Admission NIHSS may partly lie on the pathway between pre-stroke dietary exposure and subsequent disability; accordingly, Model 4 should be interpreted as an association conditional on neurological severity at presentation rather than as an estimate of the total association between DII and 90-day functional outcome. Inflammatory biomarkers and TyG were measured during the acute stroke period, and some fasting metabolic samples were obtained after initiation of acute treatment; these measures therefore should not be interpreted as direct indicators of the pre-stroke inflammatory or metabolic state. Detailed analyzer and reagent information was not retained in the research dataset. Finally, outcome assessors were not formally blinded to participants' clinical information, and the observational design cannot establish that reducing dietary inflammatory potential would improve neurological recovery. Prospective multicenter studies and intervention studies are required to address these questions.

Conclusions

Higher pre-stroke dietary inflammatory potential was associated with greater neurological deficit at admission and worse 90-day functional outcome among patients with AIS included in this single-center cohort. The association with 90-day functional outcome was observed before and after additional conditioning on admission NIHSS and was supported by analyses using an alternative binary outcome definition and restriction to patients with first-ever stroke. Exploratory ordinal analyses suggested that the association was not uniform across all mRS thresholds, whereas spline analyses did not establish a nonlinear clinical threshold. Given the potential for severity-related selection, residual confounding, and dietary measurement error, these findings should be confirmed in larger multicenter studies using prospectively standardized dietary assessment, retained energy and anthropometric data, and broader representation of stroke severity.

Abbreviations

Abbreviation

Definition

AIS

acute ischemic stroke

AP

attributable proportion

AUC

area under the curve

CI

confidence interval

DII

Dietary Inflammatory Index

E-DII
energy-adjusted Dietary Inflammatory Index
FBG
fasting blood glucose
FDR
false discovery rate
hs-CRP
high-sensitivity C-reactive protein
mRS
modified Rankin Scale
NIHSS
National Institutes of Health Stroke Scale
NLR
neutrophil-to-lymphocyte ratio
OR
odds ratio
PLR
platelet-to-lymphocyte ratio
RCS
restricted cubic spline
RERI
relative excess risk due to interaction
ROC
receiver operating characteristic
SD
standard deviation
SII
systemic immune-inflammation index
SIRI
systemic inflammation response index
SQFFQ
semi-quantitative food-frequency questionnaire
TG
triglycerides
TyG
triglyceride-glucose index

Declarations

Ethics approval and consent to participate

The study involved human participants, and written informed consent was obtained from all included participants. The protocol "Association of the Dietary Inflammatory Index and Triglyceride-Glucose Index with the Severity and Prognosis of Acute Ischemic Stroke" was approved by the Ethics Committee of the Second Affiliated Hospital of Harbin Medical University on May 29, 2024 (approval no. YJSKY2024-292; approved study period May 2024 to May 2026). Recruitment for the present cohort began on June 1, 2024, after ethics approval, and all participants included in this analysis were recruited through May 2025. All study procedures and data collection were conducted in accordance with the approved protocol and applicable institutional requirements.

Consent for publication

Not applicable.

Competing interests

The authors declare that they have no competing interests.

Authors' information

Not applicable.

Funding

This work was supported by the Key Research and Development Program of Heilongjiang Province (grant no. 2024ZX12C15) and the National Natural Science Foundation of China (grant no. 82171328).

Author Contribution

Jinan Fang (JF) and Peijia Liu (PJL) contributed equally to this work. JF: Conceptualization, Methodology, Investigation, Formal analysis (including correlation analyses), Visualization, Writing-original draft. PJL: Methodology, Investigation, Formal analysis (including correlation analyses), Writing-original draft. Fan Wang (FW), Maoqing Wang (MW), Ying Cao (YC), Zhaojun Liu (ZL), Xuesong Sun (XS), Siyi Chen (SC), Zhaobo Zhang (ZZ), and Xingbang Zou (XZ): Investigation, Data curation, Validation. Yin Huang (YH): Conceptualization, Supervision, Funding acquisition, Project administration, Writing-review and editing. Peifang Liu (PFL): Conceptualization, Supervision, Funding acquisition, Project administration, principal investigator of the National Natural Science Foundation of China project (grant no. 82171328) and the Key Research and Development Program of Heilongjiang Province project (grant no. 2024ZX12C15), Writing-review and editing. All authors read and approved the final manuscript.

Acknowledgements

Not applicable.

Availability of data and materials

The datasets used and/or analyzed during the current study are not publicly available because they contain clinical information subject to patient-privacy and institutional restrictions, but a de-identified dataset may be available from the corresponding author on reasonable request and subject to ethics/institutional approval. The English version of the dietary and nutritional assessment questionnaire used for data collection is provided as Additional file 4.

References

1. GBD 2021 Stroke Risk Factor Collaborators. Global, regional, and national burden of stroke and its risk factors, 1990–2021: a systematic analysis for the Global Burden of Disease Study 2021. *Lancet Neurol.* 2024;23(10):973–1003. 10.1016/S1474-4422(24)00369-7.
2. Hou S, Zhang Y, Xia Y, et al. Global, regional, and national epidemiology of ischemic stroke from 1990 to 2021. *Eur J Neurol.* 2024;31(12):e16481. 10.1111/ene.16481.
3. Iadecola C, Anrather J. The immunology of stroke: from mechanisms to translation. *Nat Med.* 2011;17(7):796–808. 10.1038/nm.2399.
4. Qin C, Yang S, Chu YH, et al. Signaling pathways involved in ischemic stroke: molecular mechanisms and therapeutic interventions. *Signal Transduct Target Ther.* 2022;7:215. 10.1038/s41392-022-01064-1.
5. Shivappa N, Steck SE, Hurley TG, et al. Designing and developing a literature-derived, population-based dietary inflammatory index. *Public Health Nutr.* 2014;17(8):1689–96. 10.1017/S1368980013002115.
6. Aleksandrova K, Koelman L, Rodrigues CE. Dietary patterns and biomarkers of oxidative stress and inflammation: a systematic review of observational and intervention studies. *Redox Biol.* 2021;42:101869. 10.1016/j.redox.2021.101869.
7. Shivappa N, Godos J, Hebert JR, et al. Dietary Inflammatory Index and cardiovascular risk and mortality: a meta-analysis. *Nutrients.* 2018;10(2):200. 10.3390/nu10020200.
8. Mao Y, Weng J, Xie Q, et al. Association between dietary inflammatory index and stroke in the US population: evidence from NHANES 1999–2018. *BMC Public Health.* 2024;24:50. 10.1186/s12889-023-17556-w.
9. Huang R, Lai F, Zhao L, et al. Associations between dietary inflammatory index and stroke risk: based on NHANES 2005–2018. *Sci Rep.* 2024;14:6704. 10.1038/s41598-024-57267-9.

10. Peng M, Wang L, Xia Y, et al. High Dietary Inflammatory Index is associated with increased plaque vulnerability of carotid in patients with ischemic stroke. *Stroke*. 2020;51(10):2983–9. 10.1161/STROKEAHA.120.029035.
11. Zheng J, Hoffman KL, Chen JS, et al. Dietary inflammatory potential in relation to the gut microbiome: results from a cross-sectional study. *Br J Nutr*. 2020;124(9):931–42. 10.1017/S0007114520001853.
12. Shu Y, Wu X, Wang J, et al. Associations of Dietary Inflammatory Index with prediabetes and insulin resistance. *Front Endocrinol (Lausanne)*. 2022;13:820932. 10.3389/fendo.2022.820932.
13. Yang Y, Huang X, Wang Y, et al. The impact of triglyceride-glucose index on ischemic stroke: a systematic review and meta-analysis. *Cardiovasc Diabetol*. 2023;22:2. 10.1186/s12933-022-01732-0.
14. Bukke SPN, Pathange BBR, Nelluri KDD, et al. Association of triglyceride glucose index with clinical outcomes in ischemic stroke: a retrospective study. *BMC Neurol*. 2024;24:371. 10.1186/s12883-024-03873-z.
15. Lin SF, Hu HH, Chao HL, et al. Triglyceride-Glucose Index and intravenous thrombolysis outcomes for acute ischemic stroke: a multicenter prospective-cohort study. *Front Neurol*. 2022;13:737441. 10.3389/fneur.2022.737441.
16. Yao M, Hao Y, Wang T, et al. A review of stress-induced hyperglycaemia in the context of acute ischaemic stroke: definition, underlying mechanisms, and the status of insulin therapy. *Front Neurol*. 2023;14:1149671. 10.3389/fneur.2023.1149671.
17. Yan N, Li N, Liu W, et al. Validity and reliability of a semi-quantitative food frequency questionnaire in groups at high risk for cardiovascular diseases. *Nutr J*. 2022;21(1):63. 10.1186/s12937-022-00815-8.
18. Zhong D, Zhang S, Wu B. Interpretation of the Chinese Guidelines for Diagnosis and Treatment of Acute Ischemic Stroke 2018 [in Chinese]. *Chin J Contemp Neurol Neurosurg*. 2019;19(11):897–901. 10.3969/j.issn.1672-6731.2019.11.015.
19. Yang YX. *China Food Composition Tables, Standard Edition. Volume 1, 6th ed.* Beijing: Peking University Medical Press; 2018.
20. Cai W, Xu J, Wu X, et al. Association between triglyceride-glucose index and all-cause mortality in critically ill patients with ischemic stroke: analysis of the MIMIC-IV database. *Cardiovasc Diabetol*. 2023;22:138. 10.1186/s12933-023-01864-x.
21. von Elm E, Altman DG, Egger M, Pocock SJ, Gøtzsche PC, Vandenbroucke JP, et al. The Strengthening the Reporting of Observational Studies in Epidemiology (STROBE) statement: guidelines for reporting observational studies. *Lancet*. 2007;370(9596):1453–7. 10.1016/S0140-6736(07)61602-X.
22. Lachat C, Hawwash D, Ocke MC, Berg C, Forsum E, Hornell A, et al. Strengthening the Reporting of Observational Studies in Epidemiology - Nutritional Epidemiology (STROBE-nut): an extension of the STROBE statement. *PLoS Med*. 2016;13(6):e1002036. 10.1371/journal.pmed.1002036.

23. Fan X, Wang Z, Li M, Qin Y, Yu H, Hang D, et al. Proinflammatory diet, plasma metabolome, and the risk of ischemic stroke: evidence from the UK Biobank. J Am Heart Assoc. 2026;15(12):e046192. 10.1161/JAHA.125.046192.

24. Gao W, Yu L, Zhang Y, Jin S, Chen Z, Chen X, et al. Objective nutritional indices as an independent predictor of functional outcome after endovascular therapy for acute ischemic stroke: a cohort study in a Chinese population. Front Nutr. 2025;12:1504208. 10.3389/fnut.2025.1504208.

25. Jia B, Ma R, Zhu P, Zhang R, Li S, Jiang Z. Combined nutritional and inflammatory risk stratification for predicting functional outcomes in acute ischemic stroke. Front Nutr. 2026;13:1869783. 10.3389/fnut.2026.1869783.

Figures

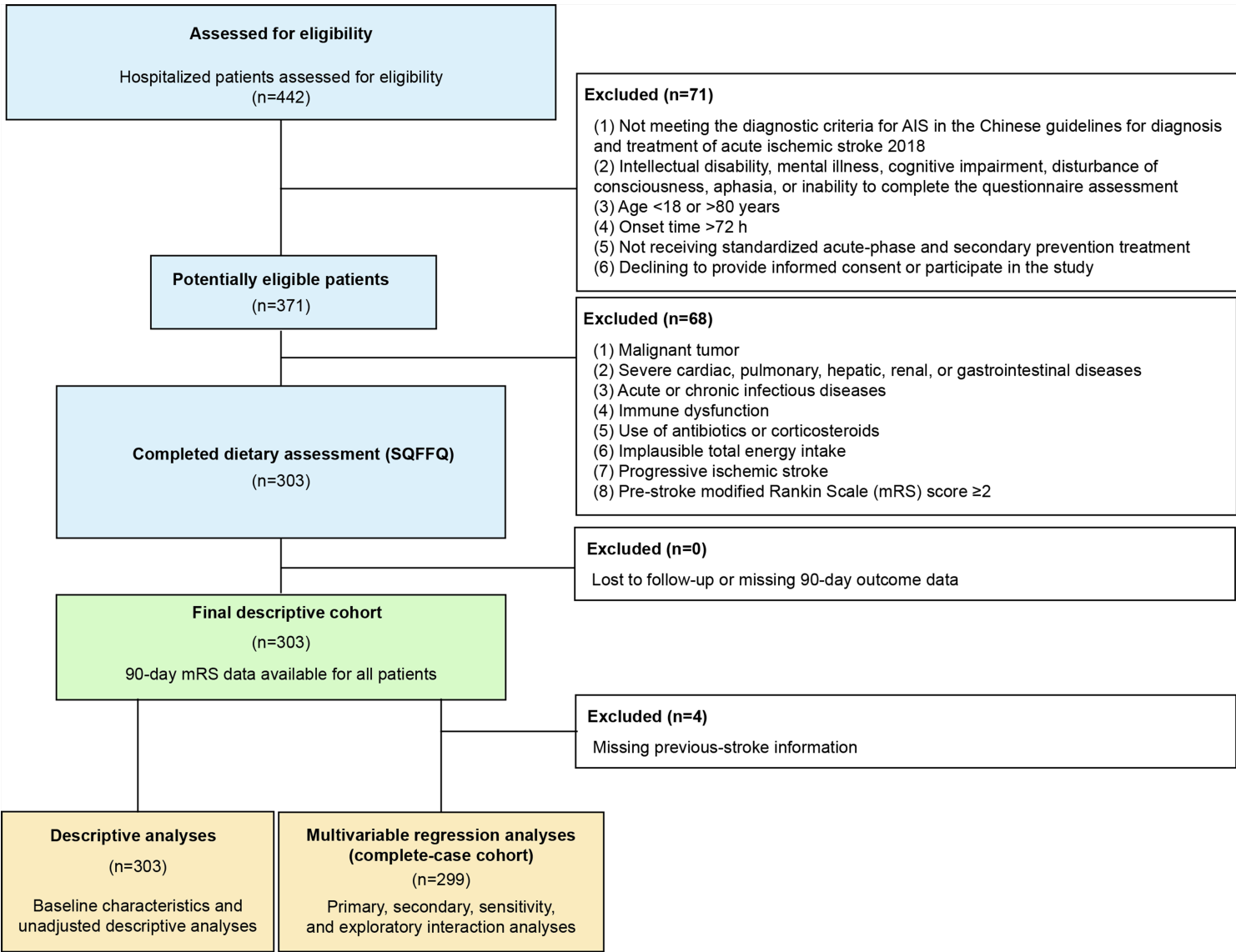


Figure 1

Participant selection and analytical cohorts

Flow diagram showing screening of 442 hospitalized patients for eligibility, two exclusion stages (71 and 68 patients, respectively), completion of dietary assessment, the final descriptive cohort of 303 patients, and the complete-case multivariable cohort of 299 patients. No participants were lost to follow-up or had missing 90-day outcome data, and 90-day mRS data were available for all 303 participants. Four participants were excluded from complete-case models because previous-stroke information was missing. Aggregate exclusion counts are shown for each screening stage together with the corresponding exclusion criteria.

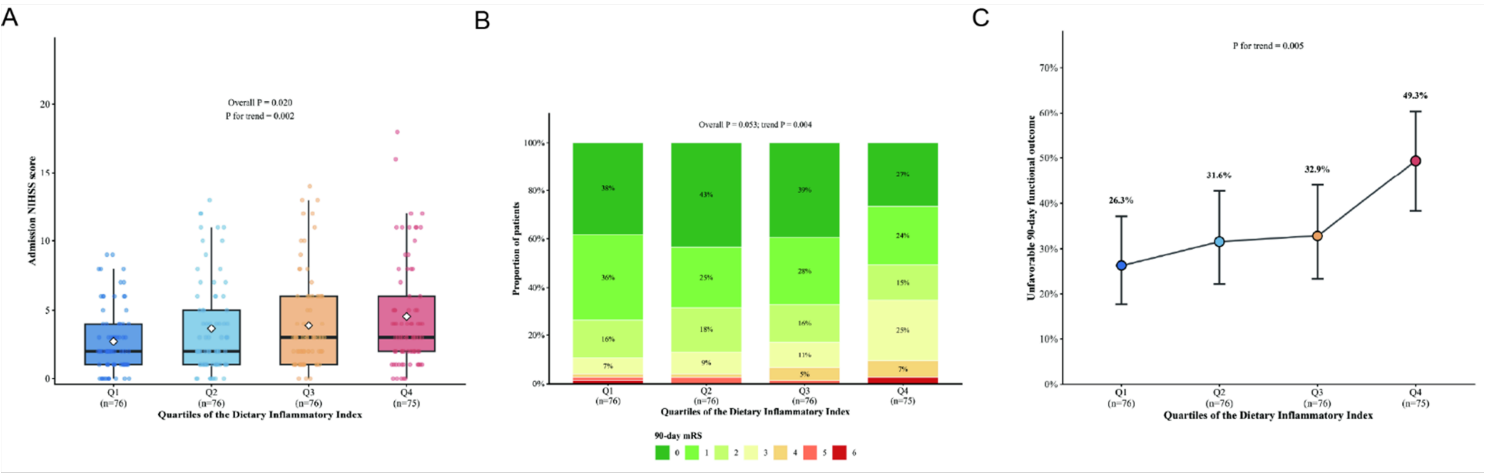


Figure 2

Stroke severity and 90-day outcomes across DII quartiles

(A) Distribution of admission National Institutes of Health Stroke Scale (NIHSS) scores across quartiles of the Dietary Inflammatory Index (DII). Boxes represent the interquartile range, horizontal lines indicate medians, whiskers extend to the most extreme observations within 1.5 times the interquartile range, points represent individual patients, and white diamonds indicate means. Overall differences were assessed using the Kruskal-Wallis test; trend was assessed using Spearman rank correlation with ordinal DII quartile. (B) Distribution of 90-day modified Rankin Scale (mRS) categories 0-6 across DII quartiles. The overall distribution was compared using Fisher's exact test with 50000 Monte Carlo replicates, and trend using Spearman rank correlation. (C) Proportion with 90-day unfavorable functional outcome (mRS 2-6 vs 0-1) across DII quartiles. Points show observed proportions and error bars show Wilson 95% CIs. P for trend was obtained from unadjusted logistic regression with quartile entered ordinally. All analyses are unadjusted. Abbreviations: DII, Dietary Inflammatory Index; mRS, modified Rankin Scale; NIHSS, National Institutes of Health Stroke Scale.

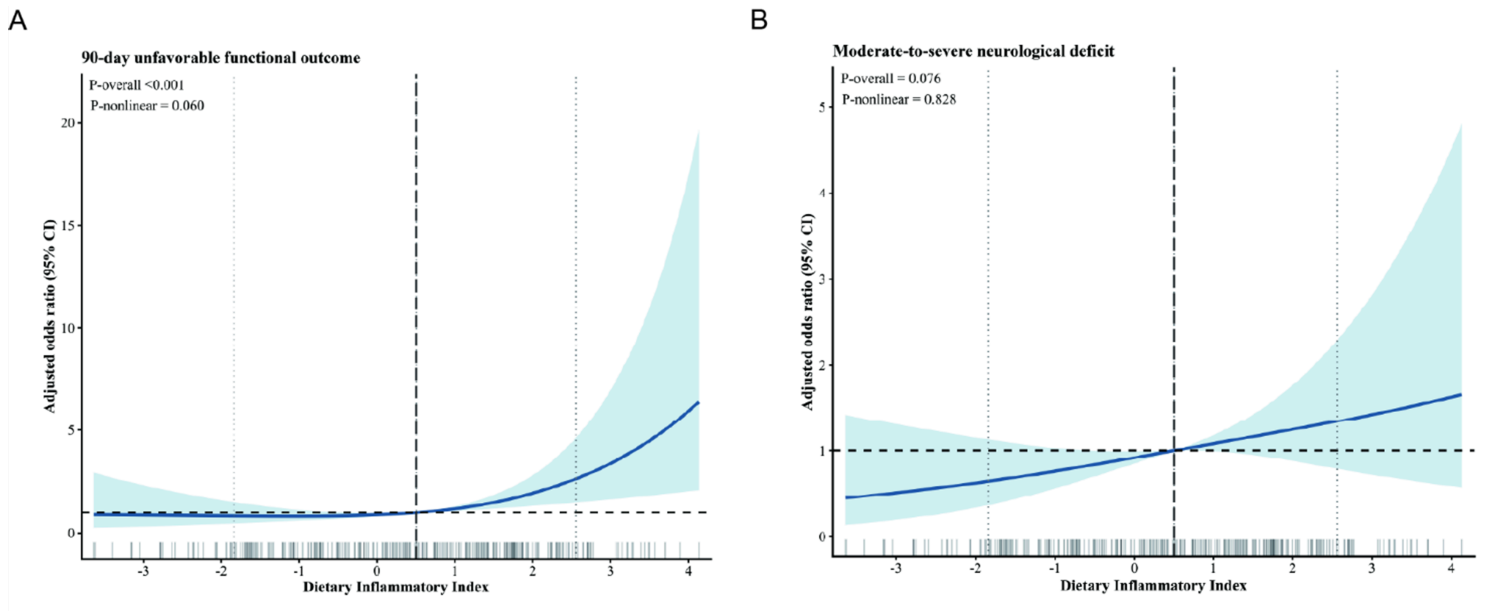


Figure 3

Dose-response associations of DII with stroke outcomes

Restricted cubic spline logistic regression models with three knots at the 10th, 50th, and 90th percentiles of DII were used for (A) 90-day unfavorable functional outcome (mRS 2-6 vs 0-1) and (B) moderate-to-severe neurological deficit at admission (NIHSS >3 vs ≤3). The median DII (0.506) was the reference. Solid curves show adjusted ORs and shaded areas show 95% CIs. Vertical dotted lines indicate knots, the vertical dash-dotted line indicates the reference, and the horizontal dashed line indicates OR=1. Rug marks show the DII distribution. Panel A was adjusted for age, sex, reported smoking history, reported alcohol drinking, hypertension, diabetes, atrial fibrillation, coronary heart disease, previous stroke, reperfusion therapy, and admission NIHSS. Panel B used the same covariates except NIHSS. P-overall compares the covariate-only and spline models; P-nonlinear compares the linear DII and spline models. Abbreviations: CI, confidence interval; DII, Dietary Inflammatory Index; mRS, modified Rankin Scale; NIHSS, National Institutes of Health Stroke

Supplementary Files

This is a list of supplementary files associated with this preprint. Click to download.

- [06Additionalfile1.docx](#)
- [08Additionalfile3.docx](#)
- [09Additionalfile4.docx](#)
- [07Additionalfile2.docx](#)