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Risk-Related Expenditure Pressure and Urban–Rural Consumption Structure in China: Provincial Evidence

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Abstract: This study examines whether provincial rural–urban differences in medical and health care expenditure shares are associated with differences in education, culture, and recreation (ECR) expenditure shares in China. It develops household shock-absorbing capacity as an analytical distinction from subjective risk preference: observed willingness to accept risk may depend not only on preferences but also on whether income, liquid resources, and social protection allow a household to withstand failure. Using a balanced panel of 31 provincial regions from 2015 to 2024, the study estimates two-way fixed-effects models with controls for provincial GDP per capita, the rural–urban disposable-income ratio, urbanization, and the share of the population aged 65 and above. In the preferred specification, a one-percentage-point increase in the rural–urban medical expenditure-share gap is associated with a 0.224-percentage-point decline in the rural–urban ECR expenditure-share gap. The coefficient remains negative when the old-age dependency ratio replaces the aging control and under both province-clustered and Driscoll–Kraay inference. However, amount-based, compositional log-ratio, first-difference, and lead-placebo specifications do not support a strong causal or absolute crowding-out interpretation. The evidence is therefore interpreted as a provincial-level association in consumption composition. Implications for demand composition and innovation are presented as bounded theoretical extensions rather than directly tested outcomes.

Keywords: medical expenditure; consumption structure; urban–rural gap; household shock-absorbing capacity; China

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1. Introduction

Household consumption is shaped not only by current income but also by the consequences of adverse events. Two households with similar observed income may respond differently to the same uncertain opportunity if one can draw on savings, credit, family support, or social insurance while the other cannot. Standard risk-preference analysis describes willingness to bear lotteries through properties of utility, most prominently the local risk-aversion measure developed by Pratt (1964). That preference-based concept is indispensable, but it does not by itself identify whether a household can finance the consequences of failure. This article uses household shock-absorbing capacity as an analytical term for that second dimension: the material and institutional ability to withstand a shock without an abrupt loss of basic consumption, debt sustainability, or longer-horizon expenditure.

The distinction matters for consumption structure. When a large part of the budget is committed to housing, health, subsistence, or other difficult-to-adjust items, an income or expenditure shock is borne disproportionately by the smaller set of adjustable purchases. Chetty and Szeidl (2007) show formally that consumption commitments can amplify aversion to moderate-stake shocks because adjustment is concentrated on flexible goods. Deaton (1991) and Carroll (1997) likewise show why liquidity constraints and income uncertainty generate buffer-stock behavior. These mechanisms suggest that observed caution can reflect constrained adjustment capacity rather than a stable psychological dislike of risk.

China provides a relevant setting because rapid income growth has coexisted with high household saving and persistent urban–rural differences. Chamon and Prasad (2010) connect rising urban saving to the increased private burden of housing, education, and health expenditure. Chamon et al. (2013) further document a sharp rise in household income uncertainty and show that uncertainty and pension reform account for more than half of the increase in urban household saving in their calibrated model. Evidence from the rollout of rural health insurance is also consistent with a consumption-allocation channel: Bai and Wu (2014) find that insurance increased nonmedical consumption, with stronger effects among poorer households and households in worse self-reported health. These studies do not imply that every increase in medical spending mechanically reduces other spending, but they show why risk protection and budget composition should be analyzed together.

The empirical question in this article is deliberately narrower than the full household mechanism. Using provincial urban and rural expenditure data, it asks whether a larger rural–urban gap in the medical and health care share of consumption is associated with a smaller rural–urban gap in the share devoted to education, culture, and recreation. The medical expenditure-share gap is treated as an imperfect proxy for relative risk-related expenditure pressure. The ECR share gap is treated as an imperfect measure of relative space for expenditure that is more adjustable, experience-oriented, or connected to education and cultural participation. Neither measure directly observes household risk preferences, household wealth buffers, or firm innovation.

The distinction between direct evidence and interpretation is essential. Provincial aggregates cannot establish that the same rural household increased medical spending and reduced ECR spending, and expenditure shares are compositional: an increase in one share must be matched by changes elsewhere. The study therefore estimates a conditional provincial-level association and subjects it to amount-based, log-ratio, lag, lead, and first-difference specifications. These checks are not presented as a search for universal significance. They are used to identify exactly which interpretation the data can and cannot sustain.

The article makes three contributions. First, it offers a conceptual contribution by separating subjective risk preference from objective shock-absorbing capacity. The latter is not proposed as a new directly observed index in this study; it is a framework that links income stability, liquid buffers, necessary expenditure commitments, and social protection to feasible risk-taking. Second, it proposes a mechanism contribution: limited shock-absorbing capacity can appear as a structural imbalance between necessary or defensive expenditure and more adjustable, longer-horizon expenditure. This reframes the demand-side question from whether total consumption rises to how the available consumption space is allocated. Third, it provides an empirical contribution based on a 2015–2024 provincial panel. The analysis documents a stable negative association within share-based fixed-effects specifications while showing that the result does not extend cleanly to absolute expenditure ratios or all alternative constructions. By separating share-based evidence from amount-based evidence, the article avoids treating a compositional association as direct proof of crowding out.

A short extension considers possible implications for demand composition and innovation. This extension is intentionally bounded. Acemoglu and Linn (2004) demonstrate in a specific pharmaceutical setting that exogenous changes in potential market size can alter innovative entry. At the same time, Mowery and Rosenberg (1979) criticize broad "demand-pull" explanations for vague definitions of demand and insufficient attention to supply-side forces. Consistent with that warning, this study does not claim to test innovation incentives. It asks only whether the documented consumption pattern could form one part of a wider demand-side mechanism that future household- and firm-level research might examine.

The remainder of the article develops the conceptual framework, describes the data and empirical strategy, reports the main and sensitivity results, and then distinguishes direct evidence from mechanism-based interpretation and downstream implications.

2. Literature review and conceptual framework

2.1. From subjective risk preference to shock-absorbing capacity

In expected-utility theory, risk aversion is conventionally defined through the curvature of utility. Pratt (1964) provides a rigorous local measure that connects utility curvature to the risk premium an individual is willing to pay to avoid a small gamble. This is a preference concept: conditional on wealth and the choice environment, it describes how the decision maker evaluates uncertain payoffs. The present study does not seek to replace that concept or redefine it as a macroeconomic indicator.

The analytical gap arises because observed risk behavior is also shaped by constraints surrounding the gamble. A household may be willing in principle to accept an uncertain investment or consumption choice but be unable to absorb a loss without missing essential payments or abandoning longer-term expenditures. Chetty and Szeidl (2007) provide a nonpsychological route from budget structure to apparently high local risk aversion. When a large fraction of expenditure is committed and costly to adjust, a moderate shock must be absorbed by a narrow set of flexible goods; the effective consumption loss in that adjustable margin becomes larger. Their unemployment-shock evidence is consistent with the predicted pattern of adjustment.

Financial fragility research makes the capacity distinction observable in a different way. Lusardi et al. (2011) measure fragility by whether households can mobilize emergency funds within a short period, regardless of whether the funds come from savings, credit, relatives, additional work, or asset sales. Their approach is useful here because it treats the ability to cope as distinct from both current income and stated attitudes toward risk. Shock-absorbing capacity in this article follows that logic. It refers to the feasible set of responses available after a shock, not to a psychological score.

This capacity depends on at least four elements: the volatility and persistence of income; liquid and collateralizable resources; the share of the budget tied to necessary or difficult-to-adjust expenditure; and institutional protection. These elements interact. A household with modest income but reliable insurance and liquid savings may have greater capacity than a higher-income household carrying rigid debt and care commitments. Conversely, a household can appear highly cautious because the downside state would force discontinuous adjustment, even if its underlying utility curvature is not unusual. This distinction is the conceptual contribution of the study, but the provincial MedicalGap variable does not directly measure it. MedicalGap is used only as a limited indicator of relative expenditure pressure that is consistent with the framework.

2.2. Uncertainty, insurance, and the allocation of consumption

The precautionary-saving literature explains why future risk affects current allocation. Deaton (1991) shows that, when borrowing is unavailable, assets operate as a buffer against unfavorable income realizations. Carroll (1997) develops the buffer-stock model further: sufficiently impatient households facing meaningful income uncertainty target a stock of wealth that protects consumption against adverse shocks. These models imply that current income alone is insufficient for understanding consumption because the expected cost of future shortfalls affects present choices.

Social insurance changes this problem by altering the household's self-insurance requirement. Hubbard et al. (1995) show that precautionary motives and means-tested social insurance can jointly reproduce important patterns of wealth accumulation; their analysis also demonstrates that the design of protection can change incentives for low-lifetime-income households. Blundell et al. (2008) distinguish the consumption response to transitory and permanent income shocks, emphasizing that households insure different shocks to different degrees. The common implication is not that risk always reduces a particular category of spending. It is that the transmission from income or expenditure shocks to consumption depends on available insurance mechanisms and household resources.

The Chinese evidence gives this mechanism institutional content. Chamon and Prasad (2010) show that urban saving increased across demographic groups and argue that rising private obligations for housing, education, and health are central to the pattern. Chamon et al. (2013) identify rising transitory income uncertainty and pension reform as major contributors to the increase in saving. These results support a broader interpretation of "capacity": when formal protection or financial intermediation is incomplete, households rely more heavily on self-insurance and may reorganize current expenditure around future contingencies.

Health insurance studies provide more direct evidence on nonmedical consumption. Exploiting the staggered introduction of China's New Cooperative Medical Scheme, Bai and Wu (2014) find that coverage increased nonmedical consumption by more than 5 percent on average and that the effect was stronger for poorer households and households in worse health. The fact that an insurance effect appeared even among households without contemporaneous out-of-pocket spending is consistent with an ex ante precautionary channel, not simply reimbursement of realized bills. The result is especially relevant to the present framework because it shows that protection against a health risk can alter the allocation of expenditure outside health care.

At the same time, medical expenditure is not a pure measure of vulnerability. Wagstaff and Lindelow (2008) show that insurance in China could increase the probability of high health expenditure by changing utilization and provider choice. A higher medical share may reflect aging, access to care, relative prices, morbidity, insurance design, or supply conditions as well as financial pressure. The empirical design therefore controls for population aging, urbanization, income differences, and economic development, and the interpretation remains explicitly proxy-based.

2.3. Necessary expenditure and development-oriented consumption space

The mechanism proposed here concerns the structure of adjustment. Medical care is not uniformly nondiscretionary, and education, culture, and recreation are not uniformly optional. Nevertheless, at the level of broad expenditure categories, unexpected or difficult-to-postpone medical needs can occupy budget space that is less flexible than many cultural, recreational, or training-related purchases. Chetty and Szeidl's commitment mechanism implies that, when rigid expenditure rises, shocks are concentrated on the remaining adjustable margin. In a constrained household, that margin may include activities whose benefits are delayed, experiential, or uncertain.

The article uses "development-oriented consumption space" as an analytical shorthand for the ECR category, not as an official statistical classification and not as a claim that all ECR expenditure raises productivity. The category combines education with diverse cultural and recreational services. Some of these purchases may be more adjustable or longer-horizon than basic consumption, but others may be committed or necessary. ECRGap is therefore retained as the empirical variable name, and the broader interpretation is treated as an incomplete proxy rather than a direct measurement.

The proposed aggregate pattern is thus a structural imbalance between a relative medical expenditure burden and the relative ECR share. This is not equivalent to a household-level statement that each yuan of medical spending displaces a yuan of education or recreation. It is a prediction about co-movement in provincial urban–rural consumption composition after observable economic and demographic differences are controlled.

2.4. Hypothesis and the bounded innovation extension

The direct empirical hypothesis is:

H1. A larger rural–urban medical expenditure-share gap is associated with a lower rural–urban ECR expenditure-share gap at the provincial level.

The hypothesis is intentionally associational. It does not state that medical expenditure causes a decline in ECR expenditure, that rural households are more risk averse, or that the ECR category is a direct measure of innovation demand.

A downstream innovation implication can nevertheless be stated as a research proposition. Firms may respond to market size, willingness to pay, adoption, and expected returns. Acemoglu and Linn (2004) use demographic shifts to identify a strong response of pharmaceutical entry and innovation to potential market size. Their design shows that demand conditions can affect innovative activity, but it is industry-specific and rests on exogenous demand variation. Mowery and Rosenberg (1979), in contrast, criticize generalized demand-pull claims when demand is vaguely defined and supply-side capabilities are not adequately separated. Taken together, these studies support only a cautious extension: if risk protection changes the composition and reliability of household demand, firms may receive different signals about which products or services can sustain experimentation. This article does not test that proposition and removes innovation from the title, dependent variables, and direct conclusions.

Table 1. Conceptual pathway and empirical status

Stage	Analytical claim	Status in this study
Subjective risk preference	Utility-based willingness to bear uncertainty	Conceptual benchmark; not measured
Shock-absorbing capacity	Ability to withstand failure using income, assets, credit, family support, and social protection	Proposed analytical concept; not directly measured
Risk-related expenditure pressure	Necessary or difficult-to-adjust expenditure can narrow the flexible budget margin	MedicalGap is an imperfect provincial proxy
Development-oriented consumption space	Part of the adjustable budget is allocated to education, culture, and recreation	ECRGap is the observed outcome proxy
Demand composition and innovation	More reliable demand may alter market signals for experimentation	Theoretical extension only; not empirically tested

Note. The table distinguishes directly observed variables from conceptual mechanisms and downstream implications.

3. Data and empirical strategy

3.1. Data sources and sample

The analysis combines annual provincial data for 31 mainland provincial-level regions from 2015 to 2024, yielding a balanced panel of 310 province–year observations. Urban and rural per capita consumption expenditure and category-level expenditure are drawn from the China Statistical Yearbook editions for 2016–2025 (National Bureau of Statistics of China, 2016–2025). Provincial GDP per capita, urban and rural disposable income, and urban and rural population are obtained from the National Bureau of Statistics provincial annual database (National Bureau of Statistics of China, n.d.). Population age structure and old-age dependency ratios are compiled from the corresponding annual population tables; the 2020 values are based on the Seventh National Population Census, whereas the other years mainly use annual population sample surveys.

The empirical unit is a province–year rural–urban gap. Comparing rural and urban shares within the same province and year places the two groups in a common regional and temporal setting, but it does not eliminate category-specific price differences, compositional constraints, or other aggregation problems. The design remains ecological: it compares provincial averages rather than following individual households. No inference is made about within-household substitution.

Registered urban unemployment is examined in a shorter 2015–2021 specification because a comparable complete series is unavailable for 2022–2024 in the downloaded database. With the preferred controls plus unemployment, the MedicalGap coefficient is −0.248 (SE = 0.108, $p = 0.028$), while the unemployment coefficient is statistically insignificant ($p = 0.711$). The same short sample without unemployment yields a MedicalGap coefficient of −0.259 ($p = 0.019$). Because unemployment reduces time coverage without materially changing the focal estimate, it is reported as a sensitivity specification rather than included in the preferred balanced-panel model.

3.2. Constructed variables and empirical roles

Medical and ECR shares are calculated separately for urban and rural residents as category expenditure divided by total per capita consumption expenditure. MedicalGap is the rural medical share minus the urban medical share, expressed in percentage points. ECRGap is defined analogously. A positive MedicalGap means that medical expenditure takes a larger share of the rural consumption budget than of the urban budget; a negative ECRGap means that the rural ECR share is lower.

The preferred controls are selected for theoretical relevance rather than mechanical significance. The logarithm of provincial GDP per capita controls for the province's changing level of economic development. IncomeGap is the natural logarithm of rural disposable income divided by urban disposable income; it captures changes in relative household resources and purchasing capacity. The income and consumption ratios are logged because $\ln(\text{rural}/\text{urban}) = \ln(\text{rural}) - \ln(\text{urban})$, so equal proportional differences are represented on a symmetric additive scale. A separate robustness specification uses the corresponding unlogged ratios to verify that the focal result is not driven by this transformation. Urbanization is the share of the resident population living in urban areas and controls for changing population composition and service access. Age65 is the percentage of the population aged 65 and above and directly addresses the possibility that medical expenditure shares rise because of demographic aging. The old-age dependency ratio is used as an alternative aging measure, but the two aging indicators are not entered together because their within-panel correlation is approximately 0.989.

Table 2. Constructed variables and empirical roles

Variable	Construction	Unit	Empirical role
ECRGap	Rural ECR expenditure share minus urban ECR expenditure share	Percentage points	Dependent variable; proxy for relative ECR consumption space
MedicalGap	Rural medical expenditure share minus urban medical expenditure share	Percentage points	Focal explanatory variable; proxy for relative risk-related expenditure pressure

Variable	Construction	Unit	Empirical role
ln GDPpc	Natural log of provincial GDP per capita	Log yuan	Economic-development control
IncomeGap	ln(rural disposable income / urban disposable income)	Log ratio	Relative household-resource control
Urbanization	Urban resident population / total resident population × 100	Percent	Population composition and access control
Age65	Population aged 65 or above / total population × 100	Percent	Preferred aging control
OldDependency	Population aged 65 or above / population aged 15–64 × 100	Percent	Alternative aging control
ConsumptionGap	ln(rural per capita consumption / urban per capita consumption)	Log ratio	Additional robustness control

Note. ECR denotes education, culture, and recreation. The share-gap variables are measured in percentage points; logged variables are clearly identified.

3.3. Baseline model

The preferred static panel specification is:

$$ECRGap_{it} = \beta MedicalGap_{it} + \gamma_1 \ln GDPpc_{it} + \gamma_2 IncomeGap_{it} + \gamma_3 Urbanization_{it} + \gamma_4 Age65_{it} + \alpha_i + \lambda_t + \varepsilon_{it} \quad (1)$$

where i indexes provincial regions and t indexes years. Province fixed effects, α_i, absorb time-invariant differences such as geography, persistent institutional conditions, and long-standing consumption patterns. Year fixed effects, λ_t, absorb shocks common to all provinces, including nationwide macroeconomic conditions and the exceptional consumption environment surrounding the COVID-19 period. The coefficient β is identified from within-province changes over time after the common year component and observed controls are removed.

The specification is static because the estimand of interest is a contemporaneous conditional association rather than persistence in ECRGap. With only ten annual periods, adding a lagged dependent variable would introduce additional short-panel estimation complications without directly identifying the proposed mechanism. A one-year lag of MedicalGap is therefore examined separately as a temporal sensitivity specification.

Because ECRGap and MedicalGap both enter in level form and are measured in percentage points, β has a level–level interpretation. A coefficient of −0.224 means that a one-percentage-point increase in MedicalGap is associated with a 0.224-percentage-point decrease in ECRGap, conditional on the included controls and fixed effects. The logarithmic transformation of some controls does not change the unit of β.

3.4. Model selection, diagnostics, and inference

The fixed-effects specification is supported by both the data structure and formal tests. Conditional on the preferred controls and year effects, a joint test strongly rejects the absence of province effects ($F(30, 265) = 9.82$, $p < 0.001$). Conditional on the preferred controls and province effects, the year effects are also jointly significant ($F(9, 265) = 3.20$, $p = 0.001$). A Breusch–Pagan (1980) Lagrange-multiplier test rejects pooled OLS in favor of a panel error structure ($\chi^2(1) = 244.22$, $p < 0.001$). A Hausman comparison that includes all five substantive regressors and the common year controls rejects the random-effects restrictions at the 5 percent level ($\chi^2(5) = 12.21$, $p = 0.032$). Fixed effects are therefore preferred on both statistical and substantive grounds, because persistent provincial characteristics are plausibly correlated with income, urbanization, aging, and consumption composition (Hausman, 1978).

Diagnostic tests indicate heteroskedasticity, serial correlation, and mild cross-sectional dependence. For the preferred specification, a Breusch–Pagan LM test rejects homoskedasticity (LM

= 85.37, $p < 0.001$; Breusch & Pagan, 1979). A Wooldridge-type test rejects the null of no first-order serial correlation ($F = 21.36$, $p < 0.001$; Drukker, 2003). The Pesaran CD statistic is −2.16 ($p = 0.031$), with an average pairwise residual correlation of −0.032 (Pesaran, 2015). The preferred tables therefore report standard errors clustered by province, allowing for heteroskedasticity and arbitrary within-province dependence. Driscoll–Kraay standard errors using a Bartlett kernel and one lag are reported as a sensitivity check because they are robust to broad forms of temporal and cross-sectional dependence, while their asymptotic justification is stronger for longer time dimensions than $T = 10$ (Driscoll & Kraay, 1998). Within-fixed-effects variance-inflation factors for the preferred regressors range from 1.07 to 1.24, providing no indication of material multicollinearity.

Panel unit-root tests are reported in Appendix B because $T = 10$ provides limited power. Zero-lag Fisher–ADF tests reject a unit root for the two focal share gaps and urbanization, but not for ln GDP per capita, IncomeGap, Age65, or the rural–urban consumption ratio. Because the panel also exhibits cross-sectional dependence, these tests are treated as descriptive diagnostics rather than a binary validity condition. The baseline includes year effects, and a first-difference model provides an additional check on common trending behavior.

Residual normality is not imposed as a validity condition because consistency of the fixed-effects estimator and province-clustered inference does not require normally distributed disturbances. A conventional Ramsey RESET test is likewise not treated as decisive in a specification containing full province and year effects. Functional-form sensitivity is examined more directly through amount ratios, unlogged rural–urban ratios, compositional log ratios, lag and lead specifications, and first differences.

3.5. Alternative specifications and interpretation rules

The sensitivity analysis is designed around the reviewers' two central concerns: compositional closure and endogeneity. First, a raw-ratio control specification replaces the logged rural–urban income ratio with the corresponding ratio in levels and adds the rural–urban consumption ratio in levels. Second, the amount-ratio model replaces shares with $\ln(\text{rural}/\text{urban per capita expenditure})$ for the medical and ECR categories. Third, an Aitchison-style specification defines Other as total consumption minus medical and ECR expenditure, and constructs rural–urban differences in $\ln(\text{ECR}/\text{Other})$ and $\ln(\text{Medical}/\text{Other})$. This uses a common residual consumption aggregate and recognizes that shares lie in a simplex rather than ordinary Euclidean space (Aitchison, 1982). Fourth, a one-year lag of MedicalGap examines whether the negative association precedes the observed ECRGap. Fifth, a first-difference model relates annual changes in the two gaps and in all preferred controls. Sixth, a lead-placebo model asks whether next year's MedicalGap predicts current ECRGap; such prediction would be inconsistent with a simple one-way causal interpretation. Finally, the old-age dependency ratio replaces Age65, an unemployment control is added in the available 2015–2021 sample, and several sample restrictions examine sensitivity to municipalities, Tibet, 2020, and the initial year.

The interpretation rule is explicit: a stable coefficient in share-based fixed-effects models supports a robust conditional association in consumption composition. It does not establish absolute expenditure crowding out. A nonsignificant or sign-changing amount or log-ratio result is treated as evidence that the stronger crowding-out claim is unsupported, not as a result to be hidden.

4. Empirical results

4.1. Descriptive patterns

The panel shows a persistent, but not monotonic, rural–urban difference in consumption composition. Rural medical expenditure shares exceed urban shares in 276 of 310 province–year observations (89.0 percent). Rural ECR shares are lower in 208 observations (67.1 percent), and both conditions hold in 191 observations (61.6 percent). The mean MedicalGap is 1.75 percentage points, whereas the mean ECRGap is −0.68 percentage points. These descriptive patterns do not identify a mechanism; they establish that the two relative differences frequently coexist.

Table 3. Summary statistics

Variable	N	Mean	Std. dev.	Minimum	Maximum
ECRGap (pp)	310	−0.677	1.544	−6.115	3.668
MedicalGap (pp)	310	1.754	1.404	−1.638	6.947
ln GDP per capita	310	11.095	0.430	10.199	12.334

Variable	N	Mean	Std. dev.	Minimum	Maximum
IncomeGap	310	−0.884	0.142	−1.237	−0.568
Urbanization (%)	310	62.662	11.553	28.788	89.839
Age65 (%)	310	12.581	3.361	4.984	21.902
OldDependency (%)	310	17.902	5.116	7.010	32.090
ConsumptionGap	310	−0.662	0.149	−1.164	−0.375

Note. Gap variables are rural minus urban. ECRGap and MedicalGap are measured in percentage points.

The dispersion statistics show substantial variation across province–year observations in the focal gaps, aging, and urbanization. The controls are included because they represent plausible confounders, not because their individual coefficients must be statistically significant. IncomeGap is negative throughout because rural disposable income remains below urban disposable income, but movement toward zero represents a narrowing of the relative income gap.

4.2. Baseline fixed-effects estimates

Table 4 reports a sequence of models so that the effect of each control block is visible. Column (1) includes only MedicalGap with province and year fixed effects. Column (2) adds ln GDP per capita; column (3) adds the rural–urban disposable-income ratio; column (4) adds urbanization; and column (5), the preferred model, adds the population share aged 65 and above. The MedicalGap coefficient remains negative and statistically significant throughout. The preferred estimate is smaller in magnitude than the estimate in column (1), although the coefficient does not decline monotonically across the intermediate specifications. The added economic and demographic controls therefore absorb part, but not all, of the share association.

Table 4. Two-way fixed-effects estimates

Variable	(1)	(2)	(3)	(4)	(5) Preferred
MedicalGap	−0.254** (0.103)	−0.256** (0.105)	−0.266** (0.099)	−0.263** (0.097)	−0.224** (0.085)
ln GDP per capita		−0.939 (3.054)	−0.707 (2.900)	−1.579 (2.756)	−2.679 (2.476)
IncomeGap			−3.337 (7.903)	−5.900 (7.355)	−6.096 (6.473)
Urbanization				0.157 (0.125)	0.112 (0.122)
Age65					−0.232* (0.116)
Province fixed effects	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	310	310	310	310	310
Overall R^2	0.700	0.700	0.702	0.710	0.721

Variable	(1)	(2)	(3)	(4)	(5) Preferred
Within R^2	0.123	0.124	0.129	0.155	0.186

Note. Dependent variable: ECRGap. Province-clustered standard errors are in parentheses. ** $p < 0.05$; * $p < 0.10$. Both focal gap variables are in percentage points. Within R^2 is calculated from province-demeaned variation while retaining the year effects included in each specification.

In the preferred model, a one-percentage-point increase in MedicalGap is associated with a 0.224-percentage-point reduction in ECRGap (SE = 0.085, p = 0.013). This is a provincial within-panel estimate. It means that, when a province's rural medical share rises relative to its urban medical share beyond what is explained by national year shocks and the included controls, its rural ECR share also tends to fall relative to the urban ECR share. It does not mean that medical spending by an individual household caused an equal reduction in that household's ECR spending.

Age65 is negative and marginally significant (p = 0.056), suggesting that aging may be associated with a lower rural-relative ECR share after the other variables are controlled. Because Age65 is included primarily to control confounding and because the model is not designed to identify the causal effect of aging, this coefficient is not treated as a separate substantive result.

4.3. Model diagnostics and robust inference

Table 5. Model selection and diagnostic tests

Diagnostic	Statistic	p-value	Implication
Province effects joint F	9.818	<0.001	Province fixed effects are required
Year effects joint F	3.198	0.001	Year fixed effects are jointly significant
Breusch–Pagan LM (panel effects)	244.216	<0.001	Pooled OLS rejected
Hausman FE vs. RE	12.212	0.032	Random-effects restrictions rejected; FE preferred
Heteroskedasticity test	85.368	<0.001	Heteroskedasticity detected
Wooldridge-type serial correlation	21.363	<0.001	Within-province serial correlation detected
Pesaran CD	−2.158	0.031	Mild cross-sectional dependence detected
Maximum within-FE VIF	1.245	—	No material multicollinearity

Note. Diagnostic results refer to the preferred share specification or its corresponding residuals. The Wooldridge statistic is reported as $F(1, 30)$. The short time dimension warrants cautious interpretation of asymptotic tests.

The sign and statistical significance of the preferred coefficient do not depend on province-clustered inference alone. Under province-clustered inference, the preferred coefficient is −0.224 (SE = 0.085, p = 0.013). Driscoll–Kraay inference with a Bartlett kernel and one lag produces a standard error of 0.048 and $p < 0.001$. The latter is treated as sensitivity evidence rather than the sole basis for significance because Driscoll–Kraay asymptotics are strongest when T is large.

The overall R^2 rises sharply once province effects are included because the fixed effects absorb persistent cross-province differences. For this reason, Table 4 also reports the conventional entity-within R^2, calculated from province-demeaned variation while retaining year effects. The preferred model has an overall R^2 of 0.721 and a within R^2 of 0.186. These measures describe fit at different levels; neither is evidence of causal strength or model bias.

4.4. Alternative constructions, temporal checks, and interpretation limits

Table 6. Unified sensitivity and identification-oriented specifications

Specification	Focal construction	Coefficient	Clustered SE	p-value	Interpretation
Preferred share model	ECRGap / MedicalGap	−0.224	0.085	0.013	Negative share association
Raw-ratio controls	ECRGap / MedicalGap; income and consumption ratios in levels	−0.185	0.085	0.038	Log-ratio transformation is not decisive
Amount-ratio model	Dependent: ln(rural/urban ECR amount) Focal: ln(rural/urban medical amount)	0.050	0.124	0.690	No evidence of absolute amount crowding out
One-year lag	ECRGap / lagged MedicalGap	−0.186	0.071	0.013	Suggestive temporal ordering only
First difference	ΔECRGap / ΔMedicalGap	−0.052	0.057	0.367	Annual changes not significantly related
Lead placebo	ECRGap / next-year MedicalGap	−0.221	0.116	0.066	Persistent common factors remain plausible
CoDA log-ratio	Rural–urban gaps in ln(ECR/Other) and ln(Medical/Other)	0.072	0.115	0.539	Baseline sign not preserved under log-ratio construction
Old-age dependency replacement	ECRGap / MedicalGap	−0.228	0.085	0.012	Aging-control result stable
Unemployment control (2015–2021)	ECRGap / MedicalGap	−0.248	0.108	0.028	Focal estimate remains significant

Note. All full-sample level models use ln GDP per capita, IncomeGap, urbanization, and Age65, except the aging-replacement model, which uses Old Dependency. The raw-ratio specification replaces IncomeGap with the rural/urban disposable-income ratio and adds the rural/urban consumption ratio in levels. Level models include province and year fixed effects. The unemployment specification uses the same controls in the available 2015–2021 sample. The first-difference model uses differenced controls and year effects.

The sensitivity results draw a clear boundary around the main finding. The raw-ratio control specification yields a MedicalGap coefficient of −0.185 (SE = 0.085, p = 0.038), indicating that the focal association is not an artifact of logging the rural–urban income and consumption ratios. The one-year lag coefficient remains negative and significant, which is consistent with temporal ordering but does not establish a causal mechanism. The amount-ratio model, however, yields a small positive and statistically insignificant coefficient. Thus, the data do not show that a higher rural-relative medical expenditure amount is accompanied by a lower rural-relative ECR expenditure amount. The evidence concerns the composition of expenditure, not a demonstrated fall in absolute ECR spending.

The compositional log-ratio result is also positive and insignificant. Aitchison's (1982) framework emphasizes that shares are constrained by closure and should often be studied through ratios among components. The failure of the baseline sign to survive this transformation means that the negative share coefficient cannot be represented as a general compositional law independent of variable construction. The correct conclusion is narrower: the ordinary expenditure-share gap contains a stable negative conditional pattern, but part of that pattern may reflect the way budget shares are defined.

The first-difference estimate is negative but imprecise, indicating that year-to-year changes do not provide the same statistical evidence as the level fixed-effects association. In addition, next year's Medical Gap is associated with current ECRGap at the 10 percent level under province-clustered inference and more strongly under Driscoll–Kraay inference, although the latter estimate is interpreted cautiously because T is short. A future variable cannot cause a current outcome in the proposed direction. This result indicates persistent trends, common omitted factors, or reverse dynamics and is a direct reason not to interpret the preferred coefficient causally.

Table 7 reports the sample-restriction checks. The coefficient remains negative in every case. It is statistically significant at the 5 percent level when Tibet, municipalities and Tibet jointly, 2020, or 2015 are excluded; excluding municipalities alone yields a coefficient of −0.160 with a province-clustered p-value of 0.056. These exclusions are not used to select a favorable result. They address the distinctive administrative and urban structure of municipalities, Tibet's unusual population scale and consumption structure, the exceptional pandemic year, and possible initial-year sensitivity.

Table 7. Preferred share model under sample restrictions

Sample	N	Province clusters	MedicalGap coefficient	Clustered SE	p-value
Baseline	310	31	−0.224	0.085	0.013
Exclude municipalities	270	27	−0.160	0.080	0.056
Exclude Tibet	300	30	−0.238	0.084	0.008
Exclude municipalities and Tibet	260	26	−0.181	0.078	0.029
Exclude 2020	279	31	−0.223	0.081	0.010
Exclude 2015	279	31	−0.220	0.089	0.019

Note. All models include ln GDP per capita, IncomeGap, urbanization, Age65, province fixed effects, and year fixed effects. Exclusions are motivated by administrative structure, extreme regional characteristics, the pandemic year, and initial-year sensitivity rather than by significance selection.

5. Discussion

5.1. What the data directly show

The direct empirical result is limited but clear. Across Chinese provinces between 2015 and 2024, a larger rural-relative medical expenditure share is associated with a smaller rural-relative ECR expenditure share after controlling for provincial development, rural–urban disposable-income differences, urbanization, aging, province fixed effects, and year fixed effects. The result remains negative under an alternative aging control and most share-based sample restrictions. This is a finding about provincial consumption composition.

The direct result does not establish three stronger claims. It does not identify a household-level behavioral response, because the same household is not observed across categories. It does not establish absolute crowding out, because the amount-ratio model is insignificant. It does not establish that medical pressure is the exogenous cause of the ECR gap, because the first-difference and lead-placebo results are inconsistent with a simple one-direction causal design. These limitations are substantive findings of the sensitivity analysis, not merely generic caveats.

5.2. Mechanism-consistent interpretation: capacity and defensive allocation

The preferred interpretation is that the provincial pattern is consistent with a capacity-constrained allocation mechanism. When households have limited buffers, necessary or difficult-to-adjust expenditures make the remaining budget more sensitive to shocks. This interpretation connects the share result to Deaton's (1991) liquidity-constrained buffer, Carroll's (1997) target-wealth behavior, Chetty and Szeidl's (2007) concentration of adjustment on flexible goods, and Lusardi et al.'s (2011) capacity-to-cope concept. It also fits the Chinese evidence that uncertainty and private expenditure obligations are associated with higher saving and that health insurance can release nonmedical consumption (Chamon & Prasad, 2010; Chamon et al., 2013; Bai & Wu, 2014).

The mechanism should not be reduced to a binary distinction between "necessary" and "luxury" consumption. Medical spending includes discretionary and preventive care, while education and cultural expenditure can be highly committed. The mechanism instead concerns relative adjustability and the consequences of failure. A household may protect current medical needs, school continuity, housing payments, or debt service and adjust recreation, some forms of training, or cultural services. Which category bears adjustment depends on institutions and household circumstances. The provincial ECR category captures only one broad margin on which such adjustment may appear.

This interpretation also explains why the conceptual contribution is not simply a relabeling of risk aversion. Subjective preference describes how a decision maker values uncertain outcomes. Shock-absorbing capacity describes whether the adverse outcome can be financed without crossing a discontinuity such as debt distress, interrupted education, or loss of essential consumption. Preferences and capacity can interact, but they are analytically different. The empirical study provides aggregate evidence compatible with the capacity channel; it does not estimate a structural parameter of either preference or capacity.

5.3. Demand composition, effective demand, and innovation as an extension

The term effective demand requires particular caution. In Keynes's (1936) framework, effective demand concerns the level of aggregate expenditure consistent with output and employment decisions. The present models do not estimate that object. They estimate a relationship between two components of household consumption. The defensible extension is therefore about the composition of realized demand: total consumption can grow while a relatively large share remains tied to health and other risk-related needs, leaving less relative space for education, cultural services, and other adjustable purchases.

This composition may matter to producers, but the connection is not automatic. Acemoglu and Linn (2004) show that exogenous increases in potential pharmaceutical market size encouraged entry of new drugs and new molecular entities. Their result establishes that demand conditions can shape innovation under a design that directly observes innovative outcomes. This article has no comparable firm, patent, R&D, or product-entry variable. It therefore cannot infer that the provincial ECR pattern changed innovation incentives.

Mowery and Rosenberg's (1979) critique is especially relevant: studies that define "demand factors" too broadly risk attributing innovation to demand while folding technological opportunity, supply capabilities, regulation, and other influences into the same label. The innovation extension here is accordingly formulated as a hypothesis for future research. If household-level evidence shows that improved protection expands reliable demand for new services, training, cultural products, or technology adoption, firm-level work could then test whether such demand affects entry, product development, or R&D. The current article supplies neither that causal link nor an innovation outcome.

5.4. Policy implications

The policy implications follow from the bounded interpretation. First, health protection can affect welfare beyond reimbursed medical bills. Bai and Wu's (2014) evidence indicates that insurance coverage can increase nonmedical consumption, especially among poorer and less healthy households, and that actual experience with reimbursement affects the response. Policies that reduce unpredictable out-of-pocket exposure, improve reimbursement reliability, and limit catastrophic payment risk may therefore expand households' feasible allocation set. The present provincial results are consistent with this possibility but do not estimate the effect of a particular reform.

Second, rural–urban disposable-income differences, urbanization, and public-service availability remain relevant even though the focal coefficient survives the inclusion of the corresponding controls. Improving household capacity involves not only temporary consumption stimulus but also more stable income, access to liquid resources, and predictable public services. Such measures can reduce the cost of adverse states and may allow households to preserve longer-horizon expenditure when shocks occur.

Third, policies should not target ECR shares mechanically. A lower medical share is not necessarily desirable if it reflects unmet care, and a higher ECR share is not automatically productive. The appropriate objective is to reduce involuntary financial exposure and improve the household's capacity to choose among consumption categories. The structure of consumption is an indicator of constraints, not a policy target in isolation.

5.5. Limitations and future research

The study has five principal limitations. First, provincial aggregates create an ecological-inference problem. Household microdata are required to test whether medical shocks, insurance, assets, and ECR spending move together within the same household. Second, MedicalGap is an imperfect proxy that may reflect morbidity, utilization, prices, supply, insurance, and aging. Third, ECRGap combines heterogeneous categories and does not directly measure human-capital investment, adoption of innovative goods, or demand quality. Fourth, compositional closure materially affects the interpretation: the log-ratio and amount models do not reproduce the baseline negative result. Fifth, no exogenous policy shock or credible instrument identifies causality, and the lead-placebo result signals unresolved dynamics.

Future work should use household panels that jointly observe health shocks, insurance reimbursement, liquid assets, debt, income uncertainty, and detailed consumption. A credible staggered policy reform or threshold design could distinguish ex ante protection from realized medical expenditure. Linking such household data to local product entry, firm sales, patents, or R&D would be necessary before the innovation extension could be evaluated. The present article should be read as a disciplined provincial evidence base and conceptual mechanism, not as the completion of that research agenda.

6. Conclusion

This article examines the provincial-level relationship between rural–urban differences in medical expenditure shares and differences in education, culture, and recreation expenditure shares in China. It introduces household shock-absorbing capacity as an analytical distinction from subjective risk preference: observed caution may arise not only from utility-based preferences but also from limited resources, rigid necessary expenditure, and incomplete protection against failure.

Using a balanced panel of 31 provincial regions from 2015 to 2024, the preferred two-way fixed-effects model controls for GDP per capita, the rural–urban disposable-income ratio, urbanization, and population aging. A one-percentage-point increase in MedicalGap is associated with a 0.224-percentage-point reduction in ECRGap. The relationship remains negative when the old-age dependency ratio replaces the preferred aging control and under province-clustered and Driscoll–Kraay inference.

The stronger interpretation is not supported. The amount-ratio and compositional log-ratio coefficients are insignificant, the first-difference estimate is imprecise, and the lead-placebo result indicates unresolved common dynamics. The evidence therefore concerns a stable association in ordinary expenditure shares, not a demonstrated fall in absolute ECR expenditure and not a causal household-level crowding-out mechanism.

The principal contribution is consequently threefold: it distinguishes preference from capacity, identifies a mechanism through which constrained capacity may appear in consumption composition, and provides aggregate evidence while explicitly separating share relationships from amount relationships. Possible implications for effective demand and innovation remain theoretical extensions. Testing them requires household-level identification and direct firm or innovation outcomes.

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Appendix A. Annual urban–rural consumption-share patterns

Table A1. Annual simple averages across 31 provincial regions

Year	Urban medical (%)	Rural medical (%)	Urban ECR (%)	Rural ECR (%)	MedicalGap (pp)	ECRGap (pp)
2015	7.17	9.26	11.01	10.37	2.08	−0.63
2016	7.51	9.32	11.30	10.56	1.81	−0.74
2017	7.74	9.73	11.55	10.55	1.98	−0.99
2018	8.30	10.27	11.36	10.65	1.97	−0.71
2019	8.62	10.69	11.71	10.95	2.08	−0.76
2020	8.65	10.29	9.41	9.19	1.65	−0.22
2021	8.90	10.07	10.84	9.87	1.17	−0.97
2022	8.76	10.05	9.86	9.48	1.29	−0.39
2023	9.24	10.91	10.71	10.05	1.67	−0.66
2024	9.05	10.91	11.11	10.41	1.86	−0.69

Note. Urban and rural category values are simple averages of provincial expenditure shares. Gaps are rural minus urban.

Appendix B. Additional econometric information

Zero-lag Fisher–ADF panel tests reject the unit-root null for ECRGap ($p = 0.002$), MedicalGap ($p < 0.001$), and urbanization ($p < 0.001$), but not for ln GDP per capita, IncomeGap, Age65, or the rural–urban consumption ratio. Because each provincial series contains only ten annual observations and the panel exhibits cross-sectional dependence, these tests have limited diagnostic value and should not be interpreted mechanically. Year fixed effects and the first-difference sensitivity model provide complementary checks on common trending behavior.

In the available 2015–2021 sample, adding the registered urban unemployment rate to the preferred controls yields a MedicalGap coefficient of −0.248 (SE = 0.108, $p = 0.028$); the unemployment coefficient is −0.118 ($p = 0.711$). The same short sample without unemployment yields a MedicalGap coefficient of −0.259 (SE = 0.104, $p = 0.019$). The focal estimate is therefore not materially altered by the unemployment control, while the balanced-panel specification preserves three additional years of data.

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