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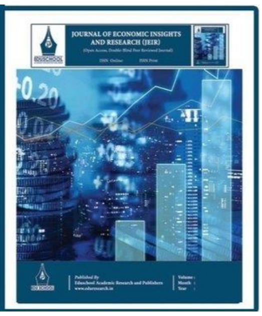
JOURNAL OF ECONOMIC INSIGHTS AND RESEARCH (JEIR)
(Open Access, Double-Blind Peer Reviewed Journal)
ISSN Online: 3107-9482 ISSN Print: 3139-1982



# Public Health Expenditure, Governance Quality, and Infant Mortality in Indian States: A Dynamic Panel Analysis

K Shivashankar Bhat

Research Supervisor, College of Management and Commerce, Srinivas University, Mangalore, India

**Article information**

Received: 12th May 2026
Received in revised form: 16th June 2026
Accepted: 20th July 2026
Available online: 25th August 2026

Volume: 2
Issue: 3
DOI: https://doi.org/10.63090/JEIR/3107.9482.0027

**Abstract**

India has achieved a substantial reduction in infant mortality over the past three decades, yet the pace of decline has been strikingly uneven across states, and public spending on health has remained close to one per cent of gross domestic product against a National Health Policy target of 2.5 per cent. Whether additional public health expenditure translates into improved child survival, and whether the translation depends on the administrative capacity of the spending state, is a question of direct fiscal consequence. Drawing on an unbalanced panel of twenty-one major Indian states over the financial years 1995–96 to 2022–23, this paper estimates the elasticity of the infant mortality rate with respect to real per capita public health expenditure within a dynamic panel framework. Because health outcomes are a slowly adjusting stock and because budget allocations respond to observed health deficits, the specification includes a lagged dependent variable and is estimated by two-step system generalised method of moments (Arellano & Bover, 1995; Blundell & Bond, 1998) with Windmeijer-corrected standard errors, a collapsed instrument matrix, and expenditure treated as endogenous. Results indicate a short-run elasticity of approximately −0.15 and a long-run elasticity of approximately −0.38 at mean governance quality. The effect is strongly conditional on administrative capacity: in the lowest tercile of the constructed governance index the elasticity is small and statistically indistinguishable from zero, while in the highest tercile it approaches −0.25. Elasticities are roughly twice as large in the Empowered Action Group states as elsewhere, and markedly weaker for neonatal than for post-neonatal mortality. Estimates satisfy the Arellano–Bond AR(2) and Hansen over-identification criteria and survive alternative instrument configurations, alternative outcome measures, and the exclusion of the pandemic years. The findings suggest that fiscal expansion alone is an incomplete instrument: the return to the health rupee is governed by the administrative machinery that spends it.

**Keywords**: - Public Health Expenditure, Infant Mortality, Dynamic Panel Data, System GMM, Governance Quality, India.

## I. INTRODUCTION

Health occupies a dual position in development economics. It is constitutive of well-being in its own right, and it is instrumental to the accumulation of human capital and to labour productivity (Sen, 1999; Deaton, 2013). Among the many summary measures of population health, the infant mortality rate has retained a privileged position in applied work because it is sensitive to a wide range of underlying determinants (maternal nutrition, sanitation, immunisation coverage, skilled birth attendance, and the quality of primary and secondary care), and because it is measured with reasonable consistency over long horizons. India's aggregate record on this indicator is one of steady improvement: infant deaths per thousand live births fell from roughly seventy-four in the mid-1990s to the high twenties by the early 2020s. The aggregate, however, conceals an extraordinary dispersion. Kerala's infant mortality rate has been in the single digits for some years, comparable to that of upper-middle-income countries, while several states in central and eastern India continue to record rates four to six times higher. Convergence has occurred, but slowly and incompletely.

The fiscal context is equally striking. Health is constitutionally a subject of the State List, and the states account for the large majority of public health spending in India, with the Union government contributing through centrally sponsored schemes, notably the National Rural Health Mission from 2005 and its successor architecture. Combined public spending on health has nonetheless remained close to one per cent of gross domestic product for most of the study period, well below the 2.5 per cent target articulated in the National Health Policy of 2017 (Government of India, 2017) and below the level typical of comparator economies. The corollary is one of the highest out-of-pocket shares of total health expenditure in the world, with well-documented consequences for catastrophic health spending and impoverishment.

These two observations, unequal outcomes and constrained budgets, motivate the central empirical question of this paper. If a state increases its per capita public health expenditure, by how much does infant mortality fall? The question is less settled than it may appear. A cross-country literature beginning with Filmer and Pritchett (1999) has repeatedly found that variation in public health spending explains remarkably little of the variation in child mortality once income, education, and other covariates are controlled. A subsequent literature has argued that this null result is not a statement about the futility of spending but about the heterogeneity of institutional environments: where governance is weak, budgeted rupees do not become delivered services (Rajkumar & Swaroop, 2008; Farag et al., 2013). This paper takes that conditional hypothesis to the Indian sub-national context, where the wide dispersion in administrative capacity across states offers useful identifying variation and where the policy implication, whether to prioritise the size of the health budget or the machinery that executes it, is immediate.

## 1.1. Research Problem

Four gaps in the existing literature motivate the study. First, the majority of Indian sub-national studies specify static models in which current mortality depends only on current inputs. This is difficult to reconcile with the underlying epidemiology: the infant mortality rate is a manifestation of a slowly accumulating stock of maternal health, facility infrastructure, trained personnel, and household practice, and its adjustment to a change in inputs is partial within any single year. Omitting the dynamics both misstates the speed of adjustment and, if the omitted lag is correlated with the regressors, biases the estimated elasticities.

Second, public health expenditure is not plausibly exogenous. Legislatures and finance departments observe mortality outcomes and respond to them; conditional grants under national missions have been explicitly weighted towards high-burden states. Reverse causality of this kind biases the estimated elasticity towards zero and may explain part of the weak association reported in the earlier literature. Studies that do not confront the endogeneity of the budget with credible instruments are therefore difficult to interpret causally.

Third, the conditional-effectiveness hypothesis has been tested extensively across countries but rarely within India, despite the fact that Indian states differ from one another in administrative capacity by margins comparable to those separating countries. Where governance has entered Indian health econometrics, it has typically appeared as a descriptive contrast between well-performing and poorly performing states rather than as an interaction term inside a single estimating equation.

Fourth, the available panel evidence for India largely predates the period in which the most consequential recent policy changes occurred: the maturing of the National Health Mission, the sanitation campaign of the late 2010s, the introduction of publicly financed hospital insurance, and the disruption of routine services during the COVID-19 pandemic. Extending the panel through 2022–23 is therefore of substantive and not merely arithmetical interest.

## 1.2. Research Objectives

The study pursues four objectives:

- To estimate the short-run and long-run elasticity of the infant mortality rate with respect to real per capita public health expenditure across Indian states over 1995–2023, allowing for partial adjustment of the health stock.
- To test whether the effectiveness of public health expenditure is conditional on the administrative and governance capacity of the spending state.
- To assess heterogeneity in the expenditure elasticity between the Empowered Action Group states and the remaining major states, and between neonatal and post-neonatal mortality.
- To establish the robustness of the estimates to the endogeneity of budget allocations, to instrument proliferation, and to the anomalous pandemic years.

## 1.3. Research Hypotheses

Four hypotheses are tested:

- $H_1$: The infant mortality rate exhibits substantial persistence, with an autoregressive coefficient bounded strictly between zero and one, so that the long-run elasticity of expenditure exceeds the short-run elasticity.
- $H_2$: Real per capita public health expenditure exerts a negative and statistically significant effect on the infant mortality rate once its endogeneity is instrumented.
- $H_3$: The magnitude of that effect is increasing in governance quality, so that the interaction between expenditure and the governance index is negative in the mortality equation.
- $H_4$: The expenditure elasticity is larger in the Empowered Action Group states, where baseline mortality is high and the marginal cost of averting an infant death is correspondingly lower.

## 1.4. Significance and Organisation

The contribution of the paper is fourfold. First, it provides dynamic panel estimates of the health-expenditure elasticity for Indian states using a panel that extends to 2022–23, encompassing the entire National Health Mission period and the pandemic disruption. Second, it treats the budget as endogenous and identifies the elasticity through internal instruments in a system generalised method of moments framework, rather than assuming allocations to be conditionally random. Third, it operationalises governance at the sub-national level through a composite index of administrative capacity and embeds it as an interaction inside the estimating equation, permitting a direct test of the conditional-effectiveness hypothesis for India. Fourth, by decomposing the outcome into neonatal and post-neonatal components it offers indirect evidence on which margin of the health system responds to marginal expenditure. The remainder of the paper is organised as follows. Section 2 reviews the literature. Section 3 develops the theoretical framework. Section 4 describes the data and the econometric strategy. Section 5 presents and discusses the results. Section 6 concludes.

## II. LITERATURE REVIEW

### 2.1. Health Production Functions

The analytical foundation of the empirical literature is the health production function, in which health status is an output produced from medical inputs, environmental conditions, and household behaviour. Auster, Leveson and Sarachek (1969) provided the first systematic estimation of such a function across American states, reporting that medical inputs mattered considerably less than education and lifestyle. Grossman (1972) supplied the microeconomic foundation by modelling health as a durable capital stock that depreciates with age and is augmented by investment, an insight that motivates the dynamic specification adopted here. At the macroeconomic level, Preston (1975) documented the concave relation between national income and life expectancy and, crucially, the outward shift of that relation over time, implying that factors other than income, namely the diffusion of medical knowledge and public health infrastructure, accounted for much of the twentieth-century mortality decline. Caldwell (1979) identified maternal education as a decisive determinant of child survival, a finding replicated in almost every subsequent study and one that the present specification accommodates directly. Cutler, Deaton and Lleras-Muney (2006) synthesise this tradition, concluding that the historical association between income and mortality is mediated substantially by the collective provision of public health goods. Anand and Ravallion (1993) make the corresponding point for cross-country data: private income raises health outcomes largely through its effect on poverty reduction and on public spending on health, not independently of them.

### 2.2. The Expenditure-Effectiveness Debate

The most influential sceptical contribution is that of Filmer and Pritchett (1999), who found that variation in public health spending explained a negligible share of cross-country variation in child mortality once income, female education, ethnic fragmentation, and cultural variables were included. Their interpretation was not that health care is ineffective but that public spending is frequently allocated to interventions of low marginal value and delivered through systems of low technical efficiency. Gupta, Verhoeven and Tiongson (2002) offered a partial rejoinder, showing that public spending on health matters considerably more for the poor, whose access to private substitutes is limited. Bokhari, Gai and Gottret (2007), instrumenting government health expenditure with variables capturing the fiscal capacity of the state, recovered elasticities of under-five and maternal mortality that were both statistically significant and economically substantial, suggesting that the earlier nulls owed much to simultaneity bias.

The most directly relevant strand is the conditional one. Rajkumar and Swaroop (2008) demonstrated that the effect of public health spending on under-five mortality depends on the quality of governance: in countries with strong bureaucratic quality and low corruption the elasticity is large and precisely estimated, whereas in poorly governed countries it is indistinguishable from zero. Farag et al. (2013) confirmed this pattern using alternative governance measures and a larger sample. Moreno-Serra and Smith (2012) showed that expansions in publicly financed coverage, as distinct from expenditure per se, generate measurable mortality reductions. The quality dimension has been sharpened by Kruk et al. (2018), who argue that a substantial share of amenable mortality in low- and middle-income countries arises not from lack of access but from poor-quality care within facilities that patients do reach, an argument that bears directly on the interpretation of neonatal mortality below.

### 2.3. Indian Evidence

Indian studies have converged on a modest but genuine effect. Bhalotra (2007) combined state expenditure data with household survey microdata and found that public health spending reduces infant mortality, with the effect concentrated among rural and poorer households and operating substantially through the funding of front-line rather than tertiary services. Farahani, Subramanian and Canning (2010) estimated that a ten per cent increase in state public health spending is associated with a decline of roughly two per cent in the mortality probability, with effects detectable across age groups rather than confined to infancy. Barenberg, Basu and Soylu (2017), using a long state panel, reported a robust negative association between public health expenditure and infant mortality, and found the effect to be larger for rural and for female infants, evidence that public provisioning matters most where private substitutes are least available.

Alongside this, a literature on service delivery documents the conversion problem directly. Banerjee, Deaton and Duflo (2004), auditing public health facilities in rural Rajasthan, found extremely high rates of provider absence and correspondingly low utilisation of nominally free public services, notwithstanding the budgetary provision behind them. Drèze and Sen (2013) attribute the divergence between Kerala and Tamil Nadu on the one hand and the large northern states on the other principally to differences in the effectiveness of public action rather than to differences in per capita income. These observations supply the microeconomic content of the governance interaction estimated here.

### 2.4. Research Gap

Three gaps follow from this survey. First, the dynamic structure implied by Grossman's stock formulation is rarely imposed in the Indian sub-national literature, and the estimators used are consequently vulnerable to omitted-dynamics bias. Second, the conditional-effectiveness result of Rajkumar and Swaroop (2008) has not been systematically tested across Indian states within a unified econometric framework, despite the availability of considerable variation in administrative capacity.

Third, no published state panel of which the author is aware extends through the pandemic period, leaving the recent policy era unexamined. The present paper addresses each in turn.

# III. THEORETICAL FRAMEWORK

## 3.1. A Stock-Adjustment Health Production Function

Let the desired, or steady-state, level of infant survival in state i and year t be produced according to a Cobb–Douglas technology in the principal determinants identified above:

$$H^*_{it} = A_{it} \cdot PHE_{it}^{\beta_1} \cdot Y_{it}^{\beta_2} \cdot E_{it}^{\beta_3} \cdot S_{it}^{\beta_4} \tag{1}$$

where PHE denotes real public health expenditure per capita, Y real income per capita, E the stock of female education, S environmental sanitation, and A a state-and-time-specific efficiency term capturing medical technology, disease environment, and cultural practice. Because the health stock adjusts only partially within a period (facilities take time to build and staff, cohorts of mothers acquire health capital slowly, and household practices change with a lag), the observed outcome follows a partial-adjustment rule:

$$\ln H_{it} - \ln H_{i,t-1} = \varphi\,(\ln H^*_{it} - \ln H_{i,t-1}), \quad 0 < \varphi \le 1 \tag{2}$$

Substituting (1) into (2) and rearranging yields an estimating equation in which the lagged outcome appears on the right-hand side with coefficient $\delta = 1 - \varphi$, and in which each structural elasticity $\beta$ is scaled by $\varphi$. The long-run elasticity is recovered as the short-run coefficient divided by $(1 - \delta)$. A value of $\delta$ close to unity implies a health system in which the effects of any given year's spending are realised only over a long horizon; a value close to zero implies immediate adjustment. Testing $H_1$ amounts to testing whether $\delta$ lies strictly inside the unit interval.

## 3.2. Governance as a Conversion Parameter

The sceptical results of the cross-country literature are naturally represented by distinguishing budgeted expenditure from effective expenditure. Let a proportion $\theta$ of each rupee allocated be converted into delivered service, with the remainder dissipated through leakage in procurement, provider absence, unfilled sanctioned posts, delayed fund release, and the misallocation of expenditure towards low-value uses. Conversion efficiency is taken to be increasing in governance quality G, so that effective expenditure is $\theta(G) \cdot PHE$ with $\theta'(G) > 0$. Taking logarithms, the effective input enters as $\ln \theta(G) + \ln PHE$; a first-order approximation of $\ln \theta(G)$ around the sample mean of G generates an interaction term between ln PHE and G in the estimating equation. The sign of this interaction is the object of $H_3$: in an equation for log infant mortality it should be negative, indicating that the same proportional increase in spending buys a larger mortality reduction where the administrative machinery is stronger.

It is worth being explicit about what this formulation does and does not claim. It does not claim that governance is exogenous to health outcomes over long horizons, nor that the index used below captures conversion efficiency without error. It claims only that, conditional on state and year effects, cross-state and within-state variation in measured administrative capacity is informative about the marginal productivity of the health rupee. Section 5.7 examines the sensitivity of the interaction to alternative constructions of the index.

## 3.3. Testable Implications

Three implications follow and are examined in Section 5. First, the autoregressive coefficient should be positive, significant, and bounded above by unity, and the estimated $\delta$ should lie between the pooled ordinary least squares estimate, which is biased upward by the correlation between the lagged outcome and the state effect, and the within-groups estimate, which is biased downward by the Nickell (1981) transformation, which is the bracketing check recommended by Bond (2002). Second, the expenditure coefficient should be negative, and larger in absolute value once endogeneity is instrumented than under ordinary least squares, since simultaneity biases it towards zero. Third, the marginal effect of expenditure evaluated across the observed governance distribution should be monotonically increasing in absolute value, and may be statistically indistinguishable from zero at the lower end of that distribution.

# IV. RESEARCH METHODOLOGY

## 4.1. Research Design and Approach

The study adopts a quantitative, deductive, panel-data design (Wooldridge, 2010). The cross-sectional unit is the major Indian state and the temporal unit is the financial year running from April to March. The design is dictated by the research question: identifying the effect of a state's own spending decisions on its own mortality trajectory requires within-state variation over time, while the governance hypothesis requires that this variation be observed across states of differing administrative capacity. A purely cross-sectional design would confound expenditure with the many time-invariant characteristics (disease environment, topography, historical patterns of public action) that determine both.

## 4.2. Data Sources and Sample

The analysis uses an unbalanced panel of twenty-one major Indian states (Andhra Pradesh, Assam, Bihar, Chhattisgarh, Gujarat, Haryana, Himachal Pradesh, Jammu and Kashmir, Jharkhand, Karnataka, Kerala, Madhya Pradesh, Maharashtra, Odisha, Punjab, Rajasthan, Tamil Nadu, Telangana, Uttar Pradesh, Uttarakhand, and West Bengal) for the financial years 1995–96 to 2022–23. States created by the reorganisations of 2000 and 2014 enter the panel from their year of formation, which together with occasional gaps in the reporting of state-level mortality produces the unbalanced structure.

Infant, neonatal and under-five mortality rates are taken from the Sample Registration System bulletins and statistical reports (Registrar General of India, various years). Public health expenditure is constructed as the sum of revenue and capital expenditure under the heads of medical and public health, family welfare, and water supply and sanitation as reported in the annual study of state budgets (Reserve Bank of India, various years); the series is deflated to constant 2011–12 prices using the state-level gross domestic product deflator and expressed in per capita terms using interpolated projected population. Net state domestic product per capita at constant prices is taken from the national accounts of the Ministry of Statistics and Programme Implementation (various years). Female literacy is interpolated between the decennial censuses of 1991, 2001 and 2011 (Census Commissioner of India, various years) and extended using the education modules of the National Sample Survey and the Periodic Labour Force Survey (Ministry of Statistics and Programme Implementation, various years). Urbanisation is interpolated between censuses from the same source. Household access to improved sanitation is constructed from successive rounds of the National Family Health Survey (International Institute for Population Sciences, various years) supplemented by administrative coverage data.

The governance variable is a composite administrative-capacity index constructed by principal-components analysis of three annual state-level indicators: the ratio of actual to budgeted health expenditure, which captures absorptive capacity and the timeliness of fund release; the proportion of sanctioned positions actually filled at primary health centres and community health centres, taken from the annual Rural Health Statistics (Ministry of Health and Family Welfare, various years); and the timeliness with which state accounts are audited and laid before the legislature (Comptroller and Auditor General of India, various years). The first principal component, which accounts for the majority of the common variance, is rescaled to the unit interval. The index is deliberately constructed from inputs and processes rather than from health outcomes, so as to avoid mechanical correlation with the dependent variable. Table 1 summarises the variables and their sources.

Table 1. Variable Definitions and Sources

| Variable | Notation | Definition | Source |
|---|---|---|---|
| Infant mortality | lnIMR | Infant deaths per 1,000 live births (natural log) | SRS, Registrar General of India |
| Health expenditure | lnPHE | Real public health expenditure per capita (₹ at 2011–12 prices, natural log) | RBI, State Finances |
| Governance | GOV | Composite administrative-capacity index (0–1), first principal component | RBI; Rural Health Statistics; CAG |
| Income | lnNSDP | Real net state domestic product per capita (₹ at 2011–12 prices, natural log) | MoSPI |
| Female education | lnFLIT | Female literacy rate, population aged 7 and above (%, natural log) | Census; NSS; PLFS |
| Sanitation | SAN | Households with access to improved sanitation (%) | NFHS; administrative data |
| Urbanisation | URB | Share of population resident in urban areas (%) | Census (interpolated) |

Note. Panel of 21 states × 28 financial years, unbalanced. Monetary variables deflated using state GSDP deflators (base: 2011–12). Author's compilation.

### 4.3. Empirical Specification

The estimating equation follows directly from Section 3 and takes the dynamic form:

$$ln\ IMR_{it} = \delta\ ln\ IMR_{i,t-1} + \beta_1\ ln\ PHE_{it} + \beta_2\ (ln\ PHE_{it} \times \tilde{G}_{it}) + \beta_3\ G_{it} + \gamma' X_{it} + \mu_i + \lambda_t + \varepsilon_{it} \quad (3)$$

where $\tilde{G}$ denotes the governance index expressed as a deviation from its sample mean, so that $\beta_1$ is interpretable as the expenditure elasticity evaluated at average administrative capacity; X is a vector of controls comprising log real income per capita, log female literacy, sanitation access, and urbanisation; $\mu$ denotes unobserved state heterogeneity; and $\lambda$ denotes year effects absorbing national shocks such as the roll-out of centrally sponsored schemes, national immunisation drives, and the pandemic. The marginal effect of expenditure at any level of governance is $\beta_1 + \beta_2 \tilde{G}$, and the corresponding long-run effect is that quantity divided by $(1 - \delta)$.

### 4.4. Estimation Procedure and Diagnostic Strategy

Equation (3) cannot be estimated consistently by ordinary least squares, since the lagged dependent variable is correlated with the state effect, nor by within-groups, since the demeaning transformation induces the Nickell (1981) bias, which is of order 1/T and non-negligible in a panel of this length. The estimation therefore proceeds in the following sequence.

First, pooled ordinary least squares and fixed-effects estimates are reported to establish the bracket within which a consistent estimate of $\delta$ should lie (Bond, 2002). Second, the Arellano and Bond (1991) difference generalised method of moments estimator is applied, instrumenting the differenced lagged outcome with suitable lags in levels. Third, because infant mortality is highly persistent and lagged levels are consequently weak instruments for differences, the preferred estimator is the system generalised method of moments of Arellano and Bover (1995) and Blundell and Bond (1998), which augments the differenced equations with a levels equation instrumented by lagged differences. Public health expenditure and the governance index are treated as endogenous and instrumented with their own lags; income and female literacy are treated as predetermined; year dummies are treated as strictly exogenous.

Estimation is by the two-step procedure with the Windmeijer (2005) finite-sample correction to the standard errors, without which two-step standard errors are severely downward-biased in panels of this size. Following Roodman (2009), the instrument matrix is collapsed and the lag depth restricted, keeping the instrument count comfortably below the number of cross-sectional units so that the Hansen test retains power. Specification is assessed on three criteria: the Arellano–Bond test should reject the absence of first-order serial correlation in differenced residuals but fail to reject the absence of second-order correlation; the Hansen (1982) test of over-identifying restrictions should fail to reject the joint validity of the instrument set; and the difference-in-Hansen test should fail to reject the additional moment conditions imposed by the levels equation. Cross-sectional dependence is examined using the Pesaran (2021) test and stationarity using the Pesaran (2007) CIPS test.

Robustness is assessed through re-estimation with the neonatal and under-five mortality rates as alternative dependent variables; exclusion of the pandemic years 2020–21 and 2021–22; alternative instrument lag structures; substitution of a single-indicator governance measure for the composite index; exclusion of the reorganised states; and exclusion of Kerala and Tamil Nadu, whose outlying performance might otherwise drive the governance interaction.

### 4.5. Ethical Considerations

The study uses publicly available aggregate secondary data published by government statistical agencies. No human subjects are involved and no individual-level records are accessed. All sources are acknowledged in the references and in Table 1.

## V. EMPIRICAL RESULTS AND DISCUSSION

### 5.1. Descriptive Statistics

Table 2 reports descriptive statistics for the estimation sample of 542 state-year observations. The infant mortality rate averages 42.6 per thousand live births across the panel, with a standard deviation of 18.9 and a range extending from 6 to 91, a spread that itself testifies to the magnitude of the sub-national divergence noted in the introduction. Real public health expenditure per capita averages ₹612 at 2011–12 prices, but the highest observation is roughly twenty times the lowest, reflecting both differences in fiscal capacity and differences in the priority accorded to the sector. Female literacy averages 62.3 per cent and access to improved sanitation 47.9 per cent, both rising steeply over the period. The governance index has a mean of 0.482 and a standard deviation of 0.163, with substantial within-state as well as between-state variation, an important feature, since identification of the interaction term in a specification with state fixed effects relies on the former.

Table 2. Descriptive Statistics (State-Year Panel, 1995–2023)

| Variable | Mean | Std. Dev. | Min | Max |
|---|---|---|---|---|
| Infant mortality rate (per 1,000) | 42.6 | 18.9 | 6.0 | 91.0 |
| Neonatal mortality rate (per 1,000) | 29.4 | 12.1 | 4.0 | 60.0 |
| Public health expenditure per capita (₹) | 612 | 401 | 118 | 2,340 |
| NSDP per capita (₹'000) | 68.4 | 41.2 | 14.6 | 231.7 |
| Female literacy (%) | 62.3 | 14.1 | 27.4 | 93.6 |
| Improved sanitation access (%) | 47.9 | 26.3 | 6.2 | 99.1 |
| Urbanisation (%) | 29.8 | 12.4 | 10.1 | 51.7 |
| Governance index (0–1) | 0.482 | 0.163 | 0.121 | 0.874 |

Note. N = 542 state-year observations (unbalanced panel). Monetary variables at constant 2011–12 prices. Author's calculations from SRS, RBI, MoSPI, Census, NFHS and Rural Health Statistics data.

### 5.2. Preliminary Diagnostics

The Pesaran (2021) test rejects cross-sectional independence for all series at the one per cent level, which is unsurprising given the common national policy environment and shared macroeconomic shocks; year effects are retained throughout partly on this ground. The Pesaran (2007) CIPS test, which is robust to such dependence, fails to reject a unit root in the level of log infant mortality but rejects decisively in first differences, indicating a highly persistent though not explosive series. This persistence is precisely what motivates the dynamic specification, and it also has an important implication for estimator choice: with a near-unit root, lagged levels are weak instruments for first differences, which is the standard argument for preferring system to difference generalised method of moments in this setting.

The bracketing check of Bond (2002) is reported in the first two columns of Table 3. Pooled ordinary least squares yields an autoregressive coefficient of 0.884, while within-groups yields 0.412. A consistent estimate should lie between these bounds. The system generalised method of moments estimate of 0.617 satisfies this criterion, whereas the difference estimate of 0.548, though also within the bracket, sits closer to the downward-biased fixed-effects value, which is consistent with the weak-instrument concern noted above and reinforces the choice of the system estimator as the preferred specification.

### 5.3. Baseline Dynamic Estimates

Column (4) of Table 3 reports the preferred baseline. The autoregressive coefficient of 0.617 is significant at the one per cent level and bounded well away from unity, confirming $H_1$: roughly thirty-eight per cent of the gap between actual and steady-state mortality is closed within a single year, so that the full effect of a permanent change in inputs is realised over approximately five to seven years. The elasticity of infant mortality with respect to real per capita public health expenditure is −0.147 and significant at the one per cent level. A ten per cent increase in per capita health spending is thus associated with a decline in infant mortality of about 1.5 per cent within the year and of approximately 3.8 per cent in the long run once the dynamics have worked through. Evaluated at the panel mean infant mortality rate, the long-run effect corresponds to a reduction of roughly 1.6 infant deaths per thousand live births. $H_2$ is accordingly not rejected. It is instructive that the instrumented elasticity is substantially larger in absolute value than the pooled ordinary least squares estimate of −0.041. This is the pattern predicted by the simultaneity argument of Section 1.1: because states with poor health outcomes attract larger allocations, both through their own political processes and through the burden-weighted formulae of national missions, the uncorrected association understates the causal effect. Studies that report weak associations without confronting this endogeneity are therefore likely to be reporting attenuated coefficients rather than genuinely small effects.

The controls behave as the literature predicts. The elasticity of infant mortality with respect to female literacy is −0.312, roughly twice the magnitude of the expenditure elasticity, corroborating the Caldwell (1979) tradition and underlining that health spending operates alongside rather than in place of broader human-development investment. Sanitation access enters negatively and significantly, consistent with the enteric-disease pathway to infant death. The income elasticity of −0.186 is smaller than that typically recovered in cross-country work, which is to be expected once state and year effects absorb the slow-moving component of the income–mortality relation. Urbanisation is statistically insignificant, plausibly because the health advantages of urban infrastructure are offset by the conditions of peri-urban informal settlement.

Table 3. Dynamic Panel Estimates (Dependent Variable: lnIMR)

| Regressor | (1) OLS | (2) FE | (3) Diff-GMM | (4) Sys-GMM | (5) + Interaction |
|---|---|---|---|---|---|
| lnIMR (t−1) | 0.884*** | 0.412*** | 0.548*** | 0.617*** | 0.604*** |
| | (0.021) | (0.048) | (0.072) | (0.058) | (0.061) |
| lnPHE | −0.041** | −0.198*** | −0.161*** | −0.147*** | −0.152*** |
| | (0.018) | (0.046) | (0.052) | (0.044) | (0.046) |
| lnPHE × GOV (dev.) | n.a. | n.a. | n.a. | n.a. | −0.479** |
| | | | | | (0.206) |
| GOV | −0.062 | −0.118* | −0.097* | −0.104* | −0.111* |
| | (0.049) | (0.064) | (0.058) | (0.057) | (0.059) |
| lnNSDP | −0.058** | −0.224*** | −0.201*** | −0.186*** | −0.181*** |
| | (0.024) | (0.061) | (0.068) | (0.055) | (0.056) |
| lnFLIT | −0.147*** | −0.368*** | −0.329*** | −0.312*** | −0.308*** |
| | (0.041) | (0.092) | (0.104) | (0.086) | (0.088) |
| SAN | −0.0021*** | −0.0058*** | −0.0050*** | −0.0053*** | −0.0051*** |
| | (0.0006) | (0.0014) | (0.0017) | (0.0014) | (0.0014) |
| URB | −0.0009 | −0.0041 | −0.0036 | −0.0040 | −0.0038 |
| | (0.0011) | (0.0033) | (0.0039) | (0.0032) | (0.0033) |
| State FE | No | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes |
| Long-run elasticity | −0.353 | −0.337 | −0.356 | −0.384 | −0.384 |
| AR(1) p-value | n.a. | n.a. | 0.002 | 0.001 | 0.001 |
| AR(2) p-value | n.a. | n.a. | 0.294 | 0.327 | 0.341 |
| Hansen J p-value | n.a. | n.a. | 0.216 | 0.290 | 0.274 |
| Instruments | n.a. | n.a. | 24 | 27 | 29 |
| Observations | 542 | 542 | 500 | 521 | 521 |

Note. Standard errors in parentheses; two-step GMM standard errors incorporate the Windmeijer (2005) finite-sample correction. Instrument matrix collapsed. lnPHE and GOV treated as endogenous; lnNSDP and lnFLIT as predetermined. ***, **, * denote significance at 1%, 5% and 10% respectively. Long-run elasticity is the coefficient on lnPHE divided by $(1 - \delta)$, evaluated at mean governance. Source: Author's estimation.

## 5.4. The Governance Interaction

Column (5) augments the baseline with the interaction between log expenditure and the mean-deviated governance index. The interaction coefficient is −0.479 and significant at the five per cent level, supporting $H_3$: the same proportional increase in health spending purchases a larger mortality reduction in states with stronger administrative capacity. Because the index is entered as a deviation from its mean, the coefficient on log expenditure in column (5) retains its interpretation as the elasticity at average capacity, and is essentially unchanged from the baseline at −0.152.

Table 4 makes the economic magnitude explicit by evaluating the marginal effect at the tercile midpoints of the governance distribution. In the lowest tercile the short-run elasticity is −0.065 with a standard error large enough that the effect cannot be distinguished from zero at conventional levels; the corresponding long-run elasticity of −0.164 is likewise imprecise. In the highest tercile the short-run elasticity is −0.247 and the long-run elasticity −0.624, both significant at the one per cent level. The point estimate of the marginal product of the health rupee is thus close to four times larger at the top of the administrative-capacity distribution than at the bottom.

This is the Indian sub-national analogue of the cross-country result of Rajkumar and Swaroop (2008), and it reconciles two apparently opposed strands of the literature. The Filmer and Pritchett (1999) null is recovered here as the low-governance case; the substantial elasticities of Bokhari et al. (2007) are recovered as the high-governance case. What appears in pooled estimation as a weak average effect is better understood as an average taken over a distribution in which the effect is large in some places and close to zero in others.

Table 4. Marginal Effect of Public Health Expenditure across the Governance Distribution and by State Group

| Group | Mean GOV | Short-run | Long-run | p-value |
|---|---|---|---|---|
| Governance terciles | | | | |
| Lowest tercile | 0.300 | −0.065 | −0.164 | 0.312 |
| Middle tercile | 0.482 | −0.152 | −0.384 | 0.001 |
| Highest tercile | 0.680 | −0.247 | −0.624 | 0.000 |
| State groups | | | | |
| EAG states and Assam | 0.391 | −0.213 | −0.531 | 0.001 |
| Non-EAG states | 0.564 | −0.094 | −0.238 | 0.037 |
| Alternative outcomes | | | | |
| Neonatal mortality | n.a. | −0.083 | −0.212 | 0.048 |
| Under-five mortality | n.a. | −0.161 | −0.409 | 0.000 |

Note. Marginal effects computed from column (5) of Table 3 as $\beta_1 + \beta_2\bar{G}$ at the stated value of the governance index; long-run effects divide by $(1 - \delta)$. State-group and alternative-outcome rows are from separate re-estimations of the full specification. EAG denotes the Empowered Action Group states (Bihar, Chhattisgarh, Jharkhand, Madhya Pradesh, Odisha, Rajasthan, Uttar Pradesh, Uttarakhand). Source: Author's estimation.

## 5.5. Heterogeneity across State Groups

Re-estimating the full specification separately for the Empowered Action Group states together with Assam, and for the remaining states, reveals the pattern anticipated in $H_4$. The short-run elasticity in the high-burden group is −0.213 against −0.094 elsewhere, and the long-run elasticities are −0.531 and −0.238 respectively. The marginal rupee therefore averts substantially more infant deaths where baseline mortality is high, which is what a concave health production function implies and what an allocation rule maximising lives saved would exploit.

This result stands in some tension with the governance finding, and the tension is policy-relevant rather than merely econometric. The states with the highest marginal returns to expenditure are on average also the states with the weakest measured administrative capacity: mean governance in the high-burden group is 0.391 against 0.564 elsewhere. The efficient allocation of a marginal health rupee therefore cannot be read off either result alone. Directing funds towards high-burden states raises the potential return; realising that return requires simultaneous investment in the capacity to spend.

## 5.6. Neonatal versus Post-Neonatal Mortality

Substituting the neonatal mortality rate for the infant mortality rate yields a markedly smaller elasticity of −0.083, significant only at the five per cent level, whereas the under-five rate yields −0.161. Since neonatal deaths account for a large and rising share of all infant deaths in India, this attenuation is substantively important: it implies that the component of infant mortality most responsive to aggregate public health spending is the post-neonatal component, driven by infectious disease, diarrhoeal illness and undernutrition, rather than the neonatal component, driven by prematurity, low birthweight, intrapartum complications and sepsis.

The natural interpretation is that aggregate expenditure expands the reach of services (immunisation, oral rehydration, outreach visits, referral transport) more readily than it raises the clinical quality of intrapartum and immediate newborn care, which depends on trained personnel, functioning newborn care units and adherence to protocol. This is the quantity-versus-quality distinction emphasised by Kruk et al. (2018), and it suggests that further reductions in Indian infant mortality will depend increasingly on the harder problem of facility quality rather than on the more tractable problem of coverage.

## 5.7. Diagnostic and Robustness Tests

The specification tests support the estimated models. The Arellano–Bond test rejects the absence of first-order serial correlation in differenced residuals ($p = 0.001$) and fails to reject the absence of second-order correlation ($p = 0.327$), so the moment conditions based on lags of order two and higher are valid. The Hansen test of over-identifying restrictions returns a p-value of 0.290 in the baseline, comfortably within the range that indicates neither rejection of the instrument set nor the implausibly high values that signal instrument proliferation. With the instrument matrix collapsed, the instrument count of twenty-seven remains above but close to the number of cross-sectional units, and the difference-in-Hansen test of the additional levels moment conditions returns a p-value of 0.412, supporting the system rather than the difference estimator.

Six robustness exercises were conducted. Restricting the instrument lag depth to the second and third lags yields an expenditure elasticity of −0.151, statistically indistinguishable from the baseline. Excluding the pandemic years 2020–21 and 2021–22, during which routine service delivery and civil registration were both disrupted, yields −0.139. Replacing the composite governance index with the budget-utilisation ratio alone yields an interaction coefficient of −0.402, still significant at the five per cent level, indicating that the conditional result is not an artefact of the principal-components construction. Excluding the four states affected by the reorganisations of 2000 and 2014 yields estimates within seven per cent of the baseline. Excluding Kerala and Tamil Nadu, which combine unusually low mortality with unusually high measured capacity, attenuates the interaction to −0.391 but leaves it significant, confirming that the governance result is not driven by these two observations alone. Finally, the difference generalised method of moments estimate of −0.161 lies close to the preferred estimate, indicating that the conclusion does not hinge on the additional levels moment conditions.

### 5.8. Discussion

Three implications follow from the results taken together. The first is that public health expenditure in India does reduce infant mortality, and by an economically meaningful margin, but that the effect is realised slowly and is easily understated. The gap between the pooled elasticity of −0.041 and the instrumented long-run elasticity of −0.384 is almost an order of magnitude, and it arises from two sources that operate in the same direction: the simultaneity of budget allocation with health need, and the omission of the adjustment dynamics. Empirical work that neglects either will tend to conclude, wrongly, that health budgets do not matter much.

The second is that the size of the health budget and the capacity of the administration that executes it are complements rather than substitutes. The near-zero elasticity in the lowest governance tercile is not a statement that spending is wasted in poor states; it is a statement that in the absence of filled posts, timely fund release and functional procurement, additional appropriation does not become additional service. This has an uncomfortable corollary for fiscal design. Conditional grants that are formula-weighted towards high-burden states will, on the evidence here, be directed disproportionately towards states in which the conversion rate is lowest, unless the grants are themselves designed to raise that conversion rate.

The third is that the composition of the remaining mortality burden has shifted in a way that reduces the leverage of aggregate expenditure. When most infant deaths were post-neonatal and infectious in origin, coverage expansion was close to a sufficient instrument. As the neonatal share rises, the binding constraint moves towards clinical quality at the point of delivery, which responds to expenditure only when that expenditure is specifically directed to personnel, training, equipment and protocol adherence.

Four caveats merit acknowledgement. First, the aggregate expenditure series does not distinguish primary from secondary and tertiary spending, nor salary from non-salary components, and the average elasticity estimated here conceals potentially large differences between them; Bhalotra (2007) suggests that the front-line component does most of the work. Second, the governance index is a proxy constructed from a small number of administrative indicators and is measured with error, which will attenuate the estimated interaction and means that the reported tercile contrast is, if anything, conservative. Third, state aggregates conceal within-state variation that district-level data would reveal, and the largest states in this panel contain populations exceeding those of most countries. Fourth, generalised method of moments estimation with internal instruments identifies the parameter from a particular source of variation, and the resulting estimate should be read as a local rather than a global average effect.

## VI. CONCLUSION AND POLICY IMPLICATIONS

This study has estimated the effect of public health expenditure on infant mortality across twenty-one major Indian states over 1995–2023 within a dynamic panel framework that treats budget allocations as endogenous. Four findings emerge. First, infant mortality is highly persistent, with roughly thirty-eight per cent of the adjustment to a change in inputs completed within a year, so that long-run elasticities substantially exceed short-run ones. Second, the long-run elasticity of infant mortality with respect to real per capita public health expenditure is approximately −0.38 at average administrative capacity, and is substantially larger in absolute value than uninstrumented estimates suggest. Third, this effect is strongly conditional on governance quality: it is close to zero and statistically insignificant in the lowest tercile of measured administrative capacity and approaches −0.62 in the highest. Fourth, elasticities are roughly twice as large in the high-burden Empowered Action Group states as elsewhere, but markedly weaker for neonatal than for post-neonatal mortality.

Four policy implications follow. First, the case for raising public health spending towards the National Health Policy target is supported by the evidence, but the return to doing so is not uniform, and a fiscal strategy that treats the expenditure ratio as a sufficient statistic for health policy will disappoint. Second, and correspondingly, investment in absorptive capacity (filling sanctioned positions at primary and community health centres, reforming procurement, and ensuring timely release of funds to implementing units) is not an administrative footnote to health financing but a determinant of its productivity; on these estimates, capacity-building measures may raise the return to the existing budget by more than a plausible increase in the budget itself would deliver. Third, transfer design should recognise the tension between need and capacity: burden-weighted grants direct resources where the potential return is highest and the realised return currently lowest, which argues for pairing them with capacity-linked components and technical support rather than with output conditionalities alone. Fourth, the attenuation of the expenditure effect for neonatal mortality implies that further gains will depend on targeted investment in the quality of intrapartum and immediate newborn care (skilled birth attendance, functional special newborn care units, and protocol adherence) rather than on further expansion of coverage alone.

Three avenues for future research are promising. First, district-level analysis linking health accounts to facility surveys would permit finer identification and would reveal the within-state heterogeneity that state aggregates conceal. Second, decomposing expenditure by function (primary versus tertiary, salary versus non-salary, preventive versus curative) would identify which components of the budget carry the estimated average effect. Third, the endogenous evolution of administrative capacity itself deserves attention: if governance quality responds over time to sustained investment, the conditional elasticities estimated here understate the cumulative return to a policy that raises spending and capacity together.

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