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RESEARCH ARTICLE

EXAMINING THE RELATIONSHIP BETWEEN HEALTHCARE SERVICES CAPACITY AND CHILD MORTALITY IN TÜRKİYE'S NUTS-III REGIONS

Tansel KARAGÖZ *

ABSTRACT

This study examines the relationship between healthcare capacity and child mortality across Türkiye's NUTS-III regions over the period 2009–2022. To capture the multidimensional nature of local health services, a healthcare capacity index is constructed via Principal Component Analysis (PCA) using per-capita measures of health institutions, hospital beds, and healthcare personnel. The analysis then estimates panel models, with Prais–Winsten regression with panel-corrected standard errors (PCSE) as the baseline specification, and complementary fixed-effects specifications to assess robustness, while controlling for GDP per capita, female secondary school enrollment rate, and urbanization. Across specifications, higher healthcare capacity is consistently associated with lower child mortality, and higher female secondary school enrollment is also associated with lower mortality. By contrast, the estimated associations for GDP per capita and urbanization are more sensitive to model specification and should be interpreted cautiously. Moreover, although regional inequalities persist, the findings indicate that healthcare capacity has exhibited a gradual tendency toward convergence across regions over time.

Keywords: *Healthcare capacity score, child mortality rate, principal component analysis, Prais-Winsten regression*

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ARAŞTIRMA MAKALESİ

TÜRKİYE'NİN DÜZEY-III BÖLGELERİNDE SAĞLIK HİZMETLERİ KAPASİTESİ VE ÇOCUK ÖLÜM HIZI ARASINDAKİ İLİŞKİNİN İNCELENMESİ

Tansel KARAGÖZ †

ÖZ

Bu çalışma, 2009–2022 döneminde Türkiye'nin Düzey-III bölgelerindeki sağlık hizmeti kapasitesi ile çocuk ölüm oranı arasındaki ilişkiyi incelemektedir. Yerel sağlık hizmetlerinin çok boyutlu niteliğini ölçebilmek amacıyla sağlık kurumları, hastane yatakları ve sağlık personeline ilişkin kişi başına düşen göstergeler kullanılarak Temel Bileşenler Analizi (TBA) yoluyla bir sağlık kapasitesi endeksi oluşturulmuştur. Endeks ile çocuk ölüm oranı arasındaki ilişki, temel spesifikasyonda panel-düzeltilmiş standart hatalara (PCSE) dayalı Prais–Winsten tahmincisi kullanılarak analiz edilmiş; bulguların sağlamlığı ise tamamlayıcı sabit etkiler spesifikasyonlarıyla sınanmıştır. Tüm modellerde kişi başına GSYH, kadın ortaöğretim okullaşma oranı ve kentleşme oranı kontrol değişkenleri olarak yer almaktadır. Farklı spesifikasyonlar genelinde sağlık kapasitesi ve kadın ortaöğretim okullaşma seviyesinin çocuk ölüm oranıyla tutarlı ve istikrarlı biçimde negatif ilişkili olduğu görülmektedir. Buna karşılık, kişi başına GSYH ve kentleşme oranı değişkenlerine ilişkin katsayılar model seçimine daha duyarlı olup bu bulguların temkinli biçimde yorumlanması gerekmektedir. Ayrıca, bölgeler arası eşitsizlikler sürmekle birlikte, bulgular sağlık hizmeti kapasitesinin zaman içerisinde bölgeler arasında kademeli bir yakınsama eğilimi sergilediğini göstermektedir.

Anahtar Kelimeler: *Sağlık kapasite skoru, çocuk ölüm hızı, temel bileşenler analizi, Prais-Winsten regresyon*

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I. INTRODUCTION

Child and infant mortality rates have declined markedly worldwide; nevertheless, they remain central public health outcomes and are often regarded as among the most sensitive indicators of overall population health (Korkmaz et al., 2013; Pala et al., 2010). In this study, child mortality refers specifically to the under-five mortality rate, measured as the number of deaths among children under five years of age per 1,000 live births. Globally, this rate declined from 93 deaths per 1,000 live births in 1990 to 38 in 2022. Türkiye reflects this global improvement with an even more dramatic shift, falling from an estimated high of about 258 per 1,000 live births in the 1960s to 10 per 1,000 in 2022; however, this rate remains slightly higher than those observed in higher-income reference groupings, notably the OECD and the European Union, which report averages near 6 and 4 per 1,000, respectively (World Bank, 2025). Taken together, these patterns show that, alongside overall progress, substantial inequalities persist, particularly between low- and middle-income and high-income settings (Ecevit et al., 2019).

Despite Türkiye's remarkable strides in reducing child mortality over the course of more than 60 years, gaps remain in the equitable distribution of healthcare resources. Republic of Türkiye Ministry of Health (2023) reports that the child mortality rate is the highest in Southeast Anatolia (16.5 per 1,000 live births) and lowest in the West Marmara region (7.3 per 1,000 live births) in 2022, demonstrating substantial geographic inequalities in infant/child health outcomes. Understanding the root causes of child mortality and identifying effective solutions requires a localized analysis of both the factors influencing child mortality and the regional disparities in healthcare access.

A review of the literature reveals a significant gap in studies that examine the relationship between healthcare capacity and child mortality in Türkiye. This study aims to address this gap by developing a healthcare capacity index to evaluate provincial disparities and examining its relationship with child mortality across provinces in Türkiye.

1.1. Literature Review

Child mortality remains a central issue in global health, reflecting the socio-economic and healthcare disparities between and within countries. The causes of child mortality differ significantly across economic contexts. In wealthy countries, injuries and accidents are the predominant causes, while in poorer countries, the primary drivers include inadequate maternal health, malnutrition, and infectious diseases such as malaria, diarrhea, and respiratory infections (Bhalotra et al., 2008). This disparity highlights the preventable nature of most child deaths in developing countries, where targeted improvements in living conditions, healthcare access, and public services can lead to significant reductions in mortality rates.

Income levels play a foundational role in addressing child mortality. Leigh and Jencks (2007) found that higher per capita income was associated with reduced infant mortality in 12 countries. Ferrarini and Norström (2010) corroborated this relationship, showing that economic growth contributed to declining infant mortality rates in 18 high-income countries. Baird et al. (2011) highlighted the pronounced effect of GDP per capita on reducing infant mortality, particularly among female infants, across 123 countries. Li (2024) analyzes 217 countries over 2000−2022 and documents a clear inverse relationship between GDP per capita and the infant mortality rate, indicating that higher income levels are associated with fewer infant deaths. Şenol et al. (2022), using panel data for 49 middle-upper and high-income countries across 2000−2019, similarly find that higher GDP and higher per capita income are linked to lower neonatal and child mortality. Rajapakse et al. (2025) examine 19 low-income countries from 2000−2020 and report a strong negative association between per capita GDP and child mortality, underscoring the role of economic capacity in improving child survival. Gao et al. (2025) show that maternal, infant, and child mortality declined across MENA (2000−2020) and that higher GDP per capita and, even more strongly, higher current health expenditure per capita are associated with lower mortality. Tüzün (2021) similarly reports a negative association between provincial GDP per capita and both infant and child mortality, underscoring the role of regional income disparities in shaping health outcomes. Spatial analyses also reveal that provinces with lower GDP, especially in Southeastern and Eastern Anatolia, experience persistently higher infant mortality rates (Türkan et al., 2020). Tüylüoğlu and Tekin (2009) demonstrated that both per capita GDP and health expenditure significantly reduce infant mortality across 176 countries. Subaşı Ertekin et al. (2016) used cointegration analysis to demonstrate a persistent inverse relationship between per capita income and infant mortality, while noting that higher unemployment rates were associated with increased mortality. Erdoğan et al. (2013) focused on high-income OECD countries, finding a strong negative relationship between real per capita income and infant mortality rates over four decades. These findings are echoed in studies of the Turkic Republics, where Demirtaş and Metintaş (2017) identified that rising GDP per capita and health expenditures reduced infant mortality, while higher unemployment, crude birth rates, and total fertility rates amplified mortality risks.

Infrastructure, healthcare, and access to healthcare are also critical in reducing child mortality. Cesur et al. (2017) analyzed the expansion of natural gas utilization in Türkiye, showing that improved household energy sources significantly reduced infant mortality rates. Manavgat and Çelik (2017) identified income levels and social health insurance as key determinants of infant mortality in Türkiye, using spatial panel data analysis across 81 provinces. These findings align with Tatlı (2021), who found that provinces with higher literacy rates, more midwives, and greater access to health services exhibited lower infant mortality. Tüzün (2021) found that the number of hospital beds per 100.000 people and number of physicians per 100.000 people were statistically significant in two-variable models, though factors such as fertility rates, adolescent birth ratios, and GDP per capita remained critical determinants. Şimşek (2024) highlights significant disparities in healthcare capacity using normalized averages and coefficient of variation methods across Türkiye’s provinces between 2012 and 2021, emphasizing critical shortages in underdeveloped regions.

Maternal education consistently emerges as one of the most significant determinants of child mortality. A wealth of studies, including those by Asefa et al. (2000), Bicego and Boerma (1993), Shamebo et al. (1993), and Zakir and Wunnava (1999), confirm a negative relationship between maternal education and infant mortality rates. Educated mothers are better equipped to adopt health-promoting behaviors, including improved nutrition, hygiene practices, and timely use of healthcare services. This relationship is further explored by Caldwell (1979), who argued that maternal education empowers women to make informed decisions about their children’s health, even in traditional societies. Similarly, Ahmadi and Javadi (2015) found that maternal education enhances knowledge about prenatal and postnatal care, while Asefa et al. (2000) linked it to improved household hygiene and disease management. In Türkiye, Yetim et al. (2021) highlighted the significant role of university education among women in reducing infant mortality. Evirgen (2023) demonstrates that female education levels were among the most important predictors of reduced infant and child mortality rates, alongside urbanization and public health expenditure.

Urbanization has nuanced effects on child mortality. While urban areas often provide better access to healthcare and living conditions, the benefits are not universally distributed. Kimani-Murage et al. (2014) noted that disparities within urban areas, particularly in slum conditions, can negate the advantages of urbanization. Evirgen (2023) observed that higher urbanization is associated with lower child mortality in Türkiye.

This study fills a gap in the Türkiye-focused child mortality literature by moving beyond single health-input proxies and offering a consistent, province-year (NUTS-III) assessment of healthcare capacity as a multidimensional construct over 2009–2022. Its key contribution is the construction of a composite healthcare capacity score from per-capita infrastructure and workforce indicators via PCA, which allows a more integrated measurement of health-system resources and a clearer comparison of regional disparities over time. By linking this index to child mortality in a panel framework, this study adds novel evidence on how overall healthcare capacity relates to child survival and how regional inequalities in healthcare capacity evolve.

II. METHODOLOGY

To construct a comprehensive indicator of healthcare capacity from multiple correlated variables, we employ Principal Component Analysis (PCA). PCA is a widely recognized multivariate statistical technique designed to extract a lower-dimensional structure from a set of correlated variables (Bro & Smilde, 2014). It achieves this by forming linear combinations of the original variables that capture the maximal amount of variance in a sequence of uncorrelated principal components.

The healthcare capacity of a given region can be represented by multiple related indicators, such as numbers of physicians, nurses, hospital beds, and healthcare institutions, each providing overlapping facets of the overall healthcare infrastructure. As these variables are often intercorrelated, direct comparison or interpretation can be challenging. PCA reduces this complexity by summarizing the joint variation across all these indicators into fewer, orthogonal principal components. PCA here identifies a dominant pattern of variation that represents the aggregate strength of healthcare systems (Jolicoeur & Mosimann, 1960). Bro and Smilde (2014) emphasize that PCA's optimal linear combinations yield a parsimonious representation of the data, allowing us to focus on the main underlying dimension of healthcare capacity rather than disparate individual indicators.

The analysis is conducted using a province-year panel dataset covering Türkiye's 81 NUTS-III regions over 14 years from 2009 to 2022. The child mortality rate (CMR), healthcare capacity score (HCS), GDP per capita (GDPpc), female secondary school enrollment rate (FEDU), and urbanization rate (URBAN) are the core variables of the analysis. Accordingly, the panel dimensions are N = 81 provinces and T = 14 years, yielding N×T = 1,134 observations. The panel is balanced for the main analysis. The healthcare-related indicators include: number of health institutions, hospital beds, and various healthcare personnel categories, all scaled per capita. Before implementing PCA, we normalize each healthcare input variable using min-max scaling to map values into the [0,1] range, ensuring that differences in original measurement scales do not dominate the extracted components. Normalization is applied globally across the full panel, using the minimum and maximum values observed across all provinces and all years simultaneously, rather than within each year separately. This ensures that the resulting scores are comparable across time, for example, a given value of the healthcare capacity score in 2009 carries the same meaning as the same value in 2022. Formally, for each variable x, the normalized value is computed as:

$$x_i' = \frac{x_i - \min(x)}{\max(x) - \min(x)}$$

The variables used in constructing composite score are shown in Table 1.

Table 1. Variables used in PCA

Variable Name (per capita)
Health Institutions
Hospital Beds
Specialist Doctor
General Practitioners
Assistant Physicians
Dentist
Pharmacist
Nurses
Midwives
Other Healthcare Personnel

Source: TURKSTAT (2025)

Subsequently, we extract the principal components, each of which is a linear combination of the original variables. The first principal component (PC1) accounts for the maximum variance in the data, while each subsequent component (PC2, PC3, etc.) explains progressively smaller shares of variance. The components are orthogonal and therefore uncorrelated with one another.

Based on the loadings of PC1, we compute the healthcare capacity score for each region. The resulting index is subsequently normalized to range between 0 and 1 in order to ensure comparability across years.

Table 2. Eigenvalues and Explained Variance from Principal Component Analysis

Principal Component	Eigenvalue	Proportion of Variance (%)
PC1	**4.793**	**47.883**
PC2	2.404	24.016
PC3	0.808	8.076
PC4	0.658	6.578
PC5	0.460	4.593
PC6	0.290	2.902
PC7	0.242	2.417
PC8	0.150	1.496
PC9	0.116	1.158
PC10	0.088	0.881

As reported in Table 2, the first principal component has an eigenvalue of 4.793 and explains 47.9% of the total variance. The second component explains 24.0% of the variance, while each subsequent component contributes less than 10%.

Table 3. Factor Loadings for the First Principal Component (PC1)

Variable (per capita)	PC1 Loading
Nurses	0.412
Dentist	0.376
Specialist Doctor	0.366
Hospital Beds	0.356
Pharmacist	0.349
Other Healthcare Personnel	0.344
Assistant Physicians	0.312
Midwives	0.221
General Practitioners	0.167
Health Institutions	0.118

The factor loadings of the first principal component are presented in Table 3. All loadings are positive, indicating that PC1 captures a common dimension reflecting the overall level of healthcare capacity across provinces. The highest loadings are observed for nurses, dentists, specialist doctors, and hospital beds, suggesting that the index is particularly sensitive to variations in healthcare workforce and infrastructure availability.

As a first step, we estimated conventional panel specifications to diagnose unobserved heterogeneity and guide model selection. Following standard panel-data practice, we compared FE and RE using the Hausman test to evaluate whether regressors are correlated with unobserved provincial heterogeneity. In our application, this diagnostic sequence supports FE ($\chi^2(4)=153.35$, $p=0.000$) as an important robustness benchmark.

Our baseline estimator is Prais–Winsten with panel-corrected standard errors (PCSE) because the core empirical setting is time-series cross-sectional (TSCS) with many panels and relatively few time periods (here, $N > T$). In such settings, errors commonly exhibit panel heteroskedasticity, serial correlation within panels, and contemporaneous correlation across panels as in our case (see Table 6) (Beck & Katz, 1995, 1996). The Prais–Winsten transformation explicitly addresses AR(1)-type serial correlation while retaining the first observation, avoiding the loss of information associated with alternative corrections (Gujarati & Porter, 2009; Prais & Winsten, 1954). PCSEs then provide standard errors that are robust to cross-sectional dependence and heteroskedasticity, reducing the risk of overly optimistic inference when T is small relative to N (Beck & Katz, 1995, 1996).

This baseline choice is therefore driven less by a preference for a particular "structural" model (FE vs RE) and more by the error structure and finite-sample properties that are typical in short TSCS panels: PCSE is designed for precisely the case where the main threat is misspecified uncertainty due to cross-sectional dependence and heteroskedasticity with limited time variation (Beck & Katz, 1995).

To address concerns about time-invariant provincial omitted variables, we also estimate province fixed effects models. For inference, we report Driscoll–Kraay (DK) standard errors as a robustness check because they are robust to general forms of cross-sectional dependence and serial correlation as well as heteroskedasticity. That said, DK corrections are most comfortable under asymptotics with larger T; with a short panel (e.g., $T = 14$), DK remains informative but can be comparatively sensitive in finite samples. For this reason, we treat FE with Driscoll-Kraay standard errors as a robustness specification that strengthens credibility against time-invariant heterogeneity, while maintaining PCSE (Prais–Winsten) as the baseline estimator tailored to short TSCS panels where $N > T$ (Baltagi, 2008; Beck & Katz, 1995). As an additional robustness check, we also estimated a period-based specification by grouping years into broader intervals (e.g., 2009–2012, 2013–2017, 2018–2022) and including period fixed effects alongside province fixed effects. This coarser time control helps capture common shocks while preserving within-period variation in short panels. The following model is estimated:

$$\ln(CMR_{it}) = \alpha + \beta_1 HCS_{it} + \beta_2 \ln(GDPpc_{it}) + \beta_3 \ln(FEDU_{it}) + \beta_4 \ln(URBAN_{it}) + \varepsilon_{it}.$$

Here, CMR_{it} denotes the child mortality rate per 1,000 live births in province i and year t, measured as the number of deaths among children under five years of age per 1,000 live births; it enters the model in natural logarithm form as $\ln(CMR_{it})$. HCS_{it} is the healthcare capacity score and enters the model in normalized (min-max normalization) form; $GDPpc_{it}$ is GDP per capita; $FEDU_{it}$ is the female secondary school enrollment rate; $URBAN_{it}$ is the urbanization rate, each entering the model as natural logarithms; α is the intercept; β_1–β_4 are the estimated coefficients; and ε_{it} is the error term.

III. FINDINGS

This section presents the empirical findings of the study. We begin with descriptive statistics and correlation patterns to provide an overview of the distributional properties of the variables, before turning to the regression results.

Table 4. Descriptive Statistics

Variable	Observations	Mean	Std. Dev.	Minimum	Maximum
CMR (per 1,000 births)	1,134	12.77	4.28	3.00	31.80
HCS	1,134	0.42	0.17	0.00	1.00
GDPpc	1,134	8,044.29	2,999.30	2,767.00	20,883.00
FEDU (%)	1,134	79.01	15.59	20.04	100.00
URBAN (%)	1,134	74.29	18.36	31.94	100.00

Source: TURKSTAT (2025) and author's calculation

Table 4 presents descriptive statistics for the core variables. CMR exhibits considerable variation across provinces and years, ranging from 3.00 to 31.80 per 1,000 live births, with a mean of 12.77. The HCS ranges from 0 to 1 by construction, with a mean of 0.42 and moderate dispersion (SD = 0.17), reflecting meaningful differences in healthcare capacity across provinces. GDPpc varies widely, from 2,767 to 20,883, consistent with substantial regional income disparities. FEDU and URBAN also display considerable cross-provincial variation.

Table 5 presents pairwise correlations among the variables included in the regression analysis. All four explanatory variables are negatively correlated with child mortality, consistent with theoretical expectations. FEDU shows the strongest bivariate association with CMR ($r = -0.707$), followed by the HCS ($r = -0.607$) and GDPpc ($r = -0.483$), while URBAN exhibits a comparatively weaker correlation ($r = -0.251$).

Among the explanatory variables, the largest correlation is observed between HCS and FEDU ($r = 0.699$), indicating that provinces with stronger healthcare capacity also tend to have higher levels of female education. The remaining pairwise correlations range from 0.310 to 0.485. Overall, the magnitude and pattern of the correlations suggest that the variables capture related but conceptually distinct dimensions of socioeconomic development.

PCA score is obtained from the variables in Table 1. The score is a comprehensive indicator that reflects the overall healthcare capacity in each province. Higher scores indicate a better healthcare capacity, while lower scores indicate a weaker healthcare capacity. The healthcare capacity score ranks are presented in detail on the maps below.

Table 5. Correlations Among Variables

Variables	ln(CMR)	HCS	ln(GDPpc)	ln(FEDU)	ln(URBAN)
ln(CMR)	1.000				
HCS	-0.607***	1.000			
ln(GDPpc)	-0.483***	0.372***	1.000		
ln(FEDU)	-0.707***	0.699***	0.485***	1.000	
ln(URBAN)	-0.251***	0.373***	0.310***	0.382***	1.000

***p<0.01, ** p<0.05 * p<0.1

Figures 1a and 1b present provincial maps of healthcare capacity scores for 2009 and 2022 respectively, where darker shading indicates higher capacity and lighter shading indicates lower capacity. Figure 1c maps the change in rank between 2009 and 2022, with red tones reflecting relative declines and blue tones indicating improvements. Between 2009 and 2022, Türkiye's provinces underwent a visible reshuffling in the distribution of healthcare capacity. Comparing Figure 1a with Figure 1b, the overall geographic pattern remains recognizable, but it appears somewhat less rigid than before. The long-standing west-east gradient softens at the margins as several inland and eastern provinces transition into higher-capacity categories. At the same time, the broad contours of the distribution are largely preserved, with the western and northwestern provinces continuing to occupy the upper tiers and much of the east and southeast remaining in the lower ones.

The ranking changes documented in Figure 1c provide a more detailed account of where movement occurred and how pronounced it was. Kilis recorded the largest improvement in the country, rising 47 positions from 62nd to 15th place. Bayburt gained 29 positions, Konya and Kütahya each advanced 20, and Tunceli, Adıyaman, Yozgat, and Tokat all recorded gains exceeding 14 positions. Looking at the geographic profile of these provinces, a consistent pattern emerges: most of them entered the period in the lower half of the distribution and recorded some of the largest upward movements by its end, contributing to a gradual compression from below. The provinces that declined in relative standing offer an equally informative perspective on how the distribution shifted. Düzce fell 26 positions, Çankırı dropped 23, Kastamonu and Artvin each declined 20, and Muğla slipped by the same margin. Burdur, Uşak, and Amasya also recorded notable relative losses. Most of these provinces were reasonably well-positioned at the start of the period and are geographically concentrated in the Black Sea coast and parts of Western Anatolia. Their relative decline does not necessarily reflect deterioration in absolute terms; rather, it is consistent with a pattern in which lower-ranked provinces improved at a faster rate, gradually displacing those that had been more moderately positioned within the distribution. Eight provinces recorded no change in rank across the entire period: Ankara, Isparta, Edirne, Bolu, and Trabzon at the upper end of the distribution, Şanlıurfa and Şırnak near the bottom, and Karabük in an intermediate position. The stability observed at the top of the distribution reflects a durable advantage that remained intact throughout the period. At the bottom, the same stability carries a different meaning: Şanlıurfa and Şırnak remained in the lowest tiers across the full fourteen-year period, recording no meaningful shift in their relative standing despite the changes occurring elsewhere in the distribution.

Figure 1c also reveals the degree to which change has been localized rather than regionally uniform. Neighboring provinces frequently followed divergent trajectories over the same period. Kastamonu declined 21 positions while nearby Çorum improved by 13. Tunceli gained 16 positions while neighboring Elazığ moved only marginally. This within-region variation is a notable feature of the data and suggests that the process of change has been uneven at a finer geographic scale than regional categories typically capture, complicating any reading of the period as one of straightforward or uniform convergence.

Figure 1. Change in Healthcare Capacity Score (2009–2022)

Figure 1a. Healthcare Capacity Score Rank in 2009

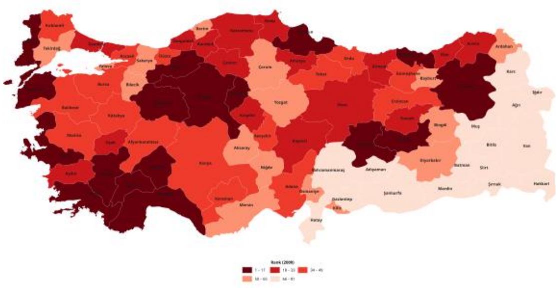


Figure 1b. Healthcare Capacity Score Rank in 2022

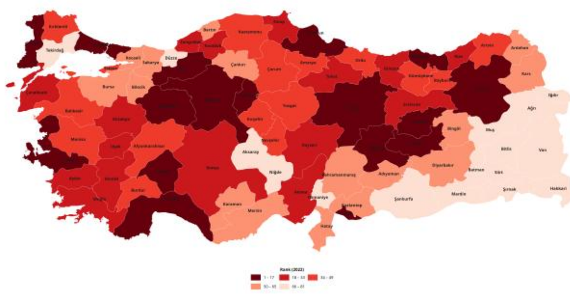
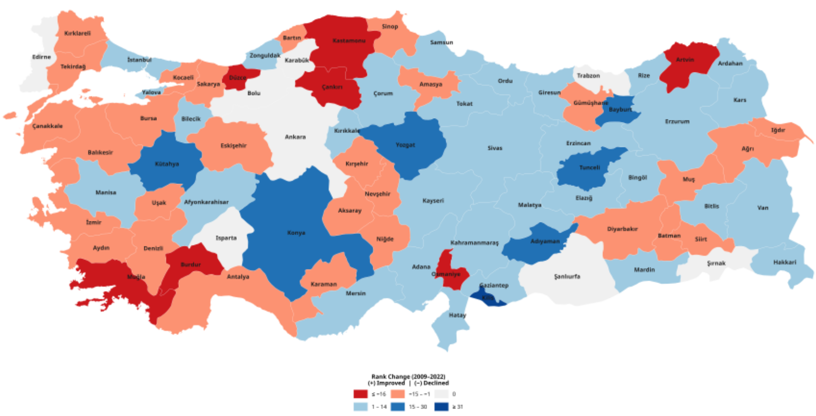
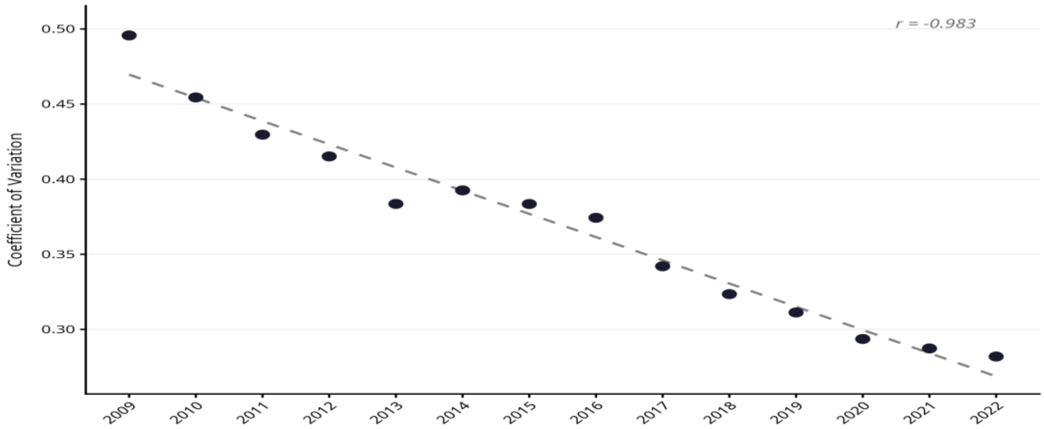


Figure 1c. Healthcare Capacity Score Rank Differences (2009–2022)



Source: Author

Figure 2. Coefficient of Variation of Healthcare Capacity Score



Source: Author

Taken together, the maps present a distribution of healthcare capacity that has shifted meaningfully over the study period, with lower-ranked provinces recording the largest relative gains and the upper end of the distribution remaining comparatively stable. Eastern and Southeastern Anatolia remains the weakest part of the distribution in both panels, and its persistence across the full period is a finding the data consistently reflects. This broader trend toward reduced spatial dispersion is further captured by the coefficient of variation of the healthcare capacity score, which declined from approximately 0.50 to 0.28 over the study period, a reduction of nearly 44 percent (see Figure 2). While this represents a substantial shift in the overall distribution, a coefficient of 0.28 still reflects considerable spatial inequality, and the picture that emerges across both the maps and the summary statistic is one of real but uneven and incomplete progress.

Table 6. Regression Diagnostics

Test	**Hypothesis**	**Statistics**	**Results**
Wooldridge / Breusch–Godfrey (Serial Correlation)	There is no first-degree autocorrelation problem.	Chi²: 26.45***	There is an autocorrelation problem.
Breusch–Pagan (Heteroskedasticity)	There is no problem of heteroskedasticity.	BP: 16.92***	There is a heteroskedasticity problem.
Pesaran CD	There is no cross-sectional dependence.	Z: 17.95***	There is a cross-sectional dependence.

***p<0.01, ** p<0.05, * p<0.1

Table 6 presents the results of the panel regression diagnostics, identifying key specification issues affecting the model's reliability. The Wooldridge Test for Serial Correlation reveals a significant problem with first-degree autocorrelation ($Chi^2 = 26.45$, $p < 0.01$). Similarly, the Breusch-Pagan test identifies heteroskedasticity, as indicated by a significant BP statistic of 16.92. Additionally, the Pesaran CD test confirms the presence of cross-sectional dependence ($Z = 17.95$, $p < 0.01$). These findings highlight the need to address autocorrelation, heteroskedasticity, and cross-sectional dependence. To ensure accurate and reliable results, standard errors were calculated using the robust Prais-Winsten approach.

Table 7. Panel Regression Results

Variables	(1) PW + PCSE	(2) FE (Provinces) + DK	(3) FE (Provinces) + Period FE + DK
HCS	-0.409*** (0.107)	-0.687*** (0.134)	-0.426** (0.185)
ln(GDPpc)	-0.153** (0.047)	0.043 (0.047)	0.037 (0.046)
ln(FEDU)	-0.614*** (0.067)	-0.336*** (0.059)	-0.133** (0.052)
ln(URBAN)	0.068 (0.045)	-0.284*** (0.081)	-0.081 (0.081)
Period: 2013–2017	—	—	-0.170*** (0.024)
Period: 2018–2022	—	—	-0.204*** (0.041)
Constant	6.417*** (0.440)	—	—
Observations	1.134	1.134	1.134
R^2	0.7437	0.3890	0.4303

***p<0.01, ** p<0.05 * p<0.1

The baseline specification is estimated using Prais-Winsten with panel-corrected standard errors (PCSE). This model exhibits a reasonable goodness-of-fit ($R^2 = 0.7437$), indicating that the included covariates explain a substantial share of the variation in log child mortality in the Time-Series Cross-Sectional (TSCS) setting. Within this baseline, healthcare capacity is negatively and statistically significantly associated with child mortality (HCS: $\beta = -0.409$, SE = 0.107, $p < 0.01$), implying that a one-unit increase in the normalized healthcare capacity score is associated with a reduction of approximately 0.409 log points in child mortality. GDP per capita is also negative and significant in the baseline (lnGDPpc: $\beta = -0.153$, SE = 0.047, $p < 0.05$), consistent with the notion that higher income levels facilitate conditions that support child survival. Female secondary education displays the largest negative association and is highly significant (lnFEDU: $\beta = -0.614$, SE = 0.067, $p < 0.01$), underscoring the central role of women's education in child health outcomes. Urbanization is positive but statistically insignificant in the baseline (lnURBAN: $\beta = 0.068$, SE = 0.045).

To assess robustness to panel alternatives and to account for error structures that may involve cross-sectional dependence and serial correlation, the model is re-estimated using a fixed-effects (within) estimator with Driscoll-Kraay standard errors (FE+DK). In this specification, the central findings remain intact: the coefficient on healthcare capacity remains negative and statistically significant (FE+DK: $\beta = -0.687$, SE = 0.134, $p < 0.01$), and female secondary education remains negative and highly significant (FE+DK: $\beta = -0.336$, SE = 0.059, $p < 0.01$). By contrast, GDP per capita becomes positive and statistically insignificant (lnGDPpc: $\beta = 0.043$, SE = 0.047), while urbanization turns negative and statistically significant (lnURBAN: $\beta = -0.284$, SE = 0.081, $p < 0.01$), suggesting that these covariates are more sensitive to whether identification relies primarily on within-province variation.

When province fixed effects are combined with period indicators (FE individual + period), healthcare capacity remains negative and statistically significant at the 5% level ($\beta = -0.426$, SE = 0.185, $p < 0.05$), and female secondary education remains negative and significant at the 5% level ($\beta = -0.133$, SE = 0.052, $p < 0.05$). In this period-FE specification, GDP per capita is close to zero and insignificant ($\beta = 0.037$, SE = 0.046), and urbanization is negative but not statistically significant ($\beta = -0.081$, SE = 0.081). The period dummies themselves are strongly negative and highly significant (2013–2017: $\beta = -0.170$, SE = 0.024, $p < 0.01$; 2018–2022: $\beta = -0.204$, SE = 0.041, $p < 0.01$), consistent with a broad downward shift in child mortality over time relative to the baseline period.

Overall, the results are most consistent for healthcare capacity and female secondary education: across PCSE, FE+DK, and FE with period fixed effects, both variables retain a negative association with child mortality and remain statistically significant, indicating a stable protective relationship across specifications. By contrast, the coefficients on GDP per capita and urbanization vary more in sign and significance across models, suggesting that their estimated associations are more sensitive to how common time shocks are absorbed and to whether identification comes primarily from within-province changes over time versus cross-province differences.

IV. DISCUSSION

Notably, the study finds a clear trend toward more balanced regional healthcare resources over time, as many historically underserved provinces in Türkiye have exhibited considerable improvements in healthcare capacity relative to some previously more advantaged areas. These gains suggest that targeted policies, such as greater public investment in health infrastructure, strategic placement of healthcare institutions, and expansions of healthcare personnel, have begun to mitigate long-standing resource gaps. In some instances, provinces once trailing in rank have vaulted forward, while others considered regionally strong have experienced relative declines, indicating a dynamic reordering of healthcare capacity across the country. Nonetheless, despite these encouraging shifts, significant differences in healthcare capacity persist, as evidenced by the continuing spread between the highest- and lowest-scoring regions.

The results clearly indicate that initiatives aimed at augmenting healthcare personnel, expanding healthcare institutions, and improving bed capacity have led to beneficial outcomes, thereby fortifying healthcare infrastructure and enhancing both the accessibility and quality of services. Such interventions have not only contributed to mitigating regional disparities but have also played a significant role in discernible reductions in child mortality. Although previous research has examined the influence of healthcare expenditures (Evirgen, 2023; Şantaş et al., 2021; Yetim et al., 2021) or individual components such as the number of healthcare personnel (Evirgen, 2023; Tatlı, 2021; Tüzün, 2021) on infant or child mortality, no prior studies have investigated the implications of an integrated healthcare capacity on child mortality rates.

Female education stands out as another central determinant. Across all model specifications, female secondary education is consistently associated with lower child mortality, underscoring its substantive importance beyond purely cross-sectional differences. The strong negative association aligns with a wide body of evidence showing that maternal schooling improves health literacy and encourages beneficial health behaviors. In particular, maternal education can foster a more health-conscious household environment, supporting better decision-making around nutrition and hygiene (Handa, 1999). Moreover, Caldwell (1979) emphasizes that when mothers, even in traditional societies, receive education, they tend to gain greater influence and control within the household. This shift in intra-household power can translate into improved decisions regarding child care, nutrition, and health inputs, thereby raising children's chances of survival and overall well-being (Caldwell, 1979). Taken together, these mechanisms suggest that women's secondary education operates not only through knowledge and practices but also through empowerment channels that reshape household priorities, reinforcing female education as a cornerstone of improved child survival and well-being.

Accordingly, the estimated associations for GDP per capita and urbanization should be interpreted with caution, as their sign and statistical significance vary across specifications in this dataset, suggesting that they are more sensitive to modeling choices than the consistently negative relationships observed for healthcare capacity and female secondary education.

V. CONCLUSION AND POLICY IMPLICATIONS

Overall, the empirical results point to two central conclusions. First, stronger healthcare capacity is associated with lower child mortality across all model specifications, indicating that regional differences in health-system resources remain closely linked to child survival. Second, female secondary education emerges as another consistent protective factor, while the estimated relationships for GDP per capita and urbanization are less stable and should therefore be interpreted more cautiously. These findings suggest that both health-system strengthening and improvements in women's education are central to reducing regional inequalities in child mortality.

Policy implications should focus on two levers that are most consistent with the evidence: strengthening healthcare capacity in lagging provinces and improving the long-run social determinants that shape child survival. In the near term, the most actionable channel is service capacity: expanding the reliability and reach of maternal-child primary care, ensuring continuity of essential inputs, and improving referral pathways so that high-risk pregnancies and pediatric emergencies are transferred quickly to higher-level facilities. Because the estimated relationship between healthcare capacity and child mortality remains negative across specifications, policies that raise effective capacity are the most defensible short-run route to further mortality reductions.

A second, very practical implication is to redesign staff assignment and retention in a way that actually makes disadvantaged regions livable and professionally attractive. Beyond simple compulsory service, this means stronger incentive bundles that restore staff welfare: higher hardship pay and housing support, childcare/schooling support for families, predictable rotations and guaranteed return rights, workload balancing through team-based staffing, protected leave, mental-health and burnout supports, and fast-tracked access to training/specialization credits. Pairing these with operational fixes such as tele-consultation links to regional hospitals, mentoring from tertiary centers, and guaranteed logistics for referrals can raise performance and retention simultaneously, which is exactly what matters if the goal is sustained capacity in low-performing provinces.

For longer-run reduction in child mortality, the stable negative association of female secondary education points to policies that keep girls in school and translate education into health-relevant capabilities. Measures that reduce dropout, increase secondary-school completion, and expand maternal health literacy (embedded in schools and community health programs) can shift care-seeking, nutrition, and hygiene practices over time. Taken together, capacity-and-workforce policies address what can change relatively quickly, while education-centered interventions anchor durable improvements both aligned with the patterns that are least sensitive to specification in the results.

These conclusions should nevertheless be interpreted in light of several limitations. One of the primary limitations is the lack of comprehensive healthcare expenditure data at the NUTS-III level in Türkiye. Although the constructed healthcare capacity score incorporates various resource-based indicators (e.g., hospital beds, healthcare personnel), the exclusion of direct expenditure measures may omit an essential aspect of resource allocation. This gap potentially restricts the study's capacity to fully capture how financial investments in health translate into outcomes such as child mortality, especially in regions where informal or external funding sources may play a more prominent role. Similarly, certain socioeconomic factors (e.g., the Gini index) could not be incorporated due to insufficient provincial data, limiting the evaluation of broader income inequality on child health outcomes.

Measurement accuracy presents another concern. In certain provinces hosting a large influx of Syrian immigrants such as Kilis, the measured "per capita" resources might be artificially inflated if refugee populations are not fully accounted for in official statistics, resulting in overestimation of healthcare capacity. In addition, the estimated relationship between healthcare capacity and child mortality may be subject to endogeneity, particularly reverse causality. Provinces with persistently high child mortality may have received targeted health investments and staffing allocations, which would raise measured healthcare capacity in response to worse outcomes. Accordingly, the coefficients should be interpreted primarily as conditional associations rather than strictly causal effects.

Future research could address these limitations by employing more extensive datasets over longer time horizons, examining additional variables pertinent to healthcare access and quality, and refining population metrics to include displaced persons.

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