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One plus one makes less than two? Consolidation policies and mortality in the Italian NHS*

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Abstract

This paper studies the population health effects of Italian Local Health Authorities' consolidation. The reform centralized administrative functions and expanded the scale of health service provision, creating entities with larger catchment areas. Using an event-study Difference-in-Differences design, we estimate the policy's impact on municipal mortality rates, accounting for heterogeneous treatment effects. Results reveal a significant increase in mortality rates starting four years after implementation, with an average 1.8% rise in total mortality observed over the following five years. Deaths from preventable conditions among individuals aged 0-74 disproportionately explain this increase. The adverse effects were primarily concentrated in municipalities within absorbed LHAs. Evidence indicates that expected economies of scale failed to improve health outcomes; instead, the reform imposed considerable health costs, particularly in municipalities belonging to larger LHAs and those with more extensive catchment area expansions. Moreover, we document that the effects were unevenly distributed, creating new vulnerable areas.

Keywords: Consolidation policy; Local healthcare units; National Health System; Mortality; Event-study
JEL classification: I11 I18 L38 G34

*In memory of Sara Pau, who passed away after a long disease at the young age of 38 before seeing the completion of this work to which she significantly contributed.

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1 Introduction

In the last decades, several countries have implemented public sector reforms to consolidate administrative functions and services provision at the local level. This process has occurred even in countries where the organization and provision of public services are decentralized. Local administrative entities and authorities have been involved in "amalgamation", "merger", or "consolidation" processes, originating larger or higher-level entities that govern larger geographical areas, with the primary goal of reducing administrative costs.

Besides evaluating cost-savings effects, numerous studies – particularly those examining municipal and county amalgamation or school district consolidations – have taken a broader perspective by assessing the impact of these reforms on welfare measures, economic growth, and other indicators linked to the institutional mission of the administrative entity. The existing evidence is mixed. Part of the literature finds that amalgamation processes reduce public expenditures and do not affect the level of services provided (e.g., Reingewertz, 2012; Ferraresi et al., 2018), increase people's subjective well-being (Wilner, 2022), may stimulate population growth, an indirect indicator of prosperity increase (Andini et al., 2017), and may improve student achievements (De Haan et al., 2016; Sandsør et al., 2021). There is, however, also evidence of adverse effects. These include exacerbating the economic divide between urban and rural areas (Hall et al., 2019), lowering returns to education and reducing educational attainments (Berry and West, 2010), as well as increasing political costs, such as reduced voter turnout (e.g., Roesel, 2017).

In this paper, we examine a distinctive form of public entity amalgamation: the consolidation of local health jurisdictions. These entities, commonly referred to as authorities, departments, or districts, are responsible for providing healthcare at the local level in many countries. Over the past two decades, examples of consolidation policies include the reduction in the number of public local health authorities in Canada and the amalgamation of Area Health Services in Australia.[1] Our analysis focuses on the merger of Local Health Authorities (LHAs) in Italy. The Italian National Health Service (NHS) is based on a decentralized structure with national, regional and local administrations. Each of the 21 autonomous regions has the authority to

[1]In Canada, the number of local authorities decreased from about 130 in 2005 to 75 in 2023 (StatisticsCanada, 2024, 2006). In Australia, a notable episode of mergers occurred in 2005 in New South Wales, where 17 Services were amalgamated into 8 larger entities (Johar and Savage, 2014). However, a legislative shift in 2011 went in the opposite direction, leading to the establishment of 15 Local Health Districts with smaller catchment areas. More recently, in 2021, in Victoria, four Services merged into a single entity.

determine the number of Local Health Authorities (LHAs) and their respective "catchment area", which include the population of the municipalities within the LHA. The LHAs are the third (local) tier of the system: they organize and provide a wide range of health services, including primary care, hospital care, outpatient clinic care, public health care, and services related to social care.

Since the 1990s, Italian national policies have aimed to curb rising healthcare costs through measures designed to enhance efficiency and contain expenditures. As part of this effort, the number of LHAs in Italy has been significantly reduced – from 228 in 1996 to 99 in 2021 – resulting in an average of approximately 611,000 enrollees per LHA, given a national population of around 60 million. This reform represents a relatively large-scale "experiment" compared to other European countries such as Germany, which currently operates approximately 400 comparable local health offices (Gesundheitsämter) for a population of about 84 million (see, e.g., RKI, 2023).[2] Catchment areas are relatively small even in the United States, with approximately 3,300 local health departments serving a population of about 335 million, although with a less directly comparable role (Cunningham et al., 2024). The catchment areas of the Italian LHAs could be comparable, in terms of average population, to the current configuration in Canada (75 LHAs for approximately 40 million inhabitants) and Australia (52 local health entities serving nearly 27 million people).

With our analysis, we aim to study whether and how this consolidation policy impacted the mission of LHAs as stated by law, which is to ensure adequate health care for individuals residing in the municipalities covered by their services. In principle, some beneficial effects could be expected if a reduction in administrative costs allows additional funds to be allocated for patient care. Furthermore, the increased scale of service provision is expected to enhance quality as healthcare professionals perform more procedures, thereby improving their skills and efficiency. However, adverse impacts on health outcomes are possible. Standardized service provision limits the adaptability of primary and outpatient care to diverse needs and varying demands. Additionally, there is a limit to the organizational complexity a single governance can manage. Finally, healthcare centralization often increases the average distance to healthcare facilities, particularly for people living in areas that are less served (typically, non-urban areas).[3]

[2]The specificity of the Italian case becomes even more evident when considering factors such as population density and urbanization. Italy has a population density of 206 inhabitants per square kilometer, compared to Germany's 240, and a lower share of its population living in urban areas (69.5% in Italy versus 76.3% in Germany). Data refer to the year 2020 and are sourced from www.Worldometers.info.

[3]Previous studies essentially focused on the relationship between population size and public health expenditure. See, in particular, Di Novi et al. (2018) for the Italian case; Santerre (2009) and Bernet and Singh (2015) for the

To assess the impact on population health of the Italian LHA amalgamation, we study the evolution of mortality at the local level. Death records represent a simple and informative measure of population health and can be very useful to assess healthcare policy interventions, especially when a sufficiently broad timespan is available.[4] We use demographic and mortality data provided by the Italian National Statistical Institute (Istat) to build mortality rates at the municipal level. We merge this information with administrative and official documentation released by the Italian Ministry of Health to reconstruct the LHAs consolidation process in the period 2002-2018. In particular, we trace the exact timing of the amalgamation and the possible transition of municipalities from one LHA to another.

The LHA consolidation reform within the Italian NHS was implemented at different times across regions, with some regions opting not to merge their LHAs, creating variation in municipal-level exposure. This staggered adoption allows us to group municipalities into twelve distinct treatment cohorts based on their year of treatment from 2002 to 2018. We estimate the causal effect on mortality rates by leveraging the variation across and within regions and in treatment exposure duration and exploiting the presence of never-treated municipalities. We employ an event-study DID approach, mainly using the Sun and Abraham (2021) estimator, which addresses heterogeneous treatment effects in time for the cohorts and captures the full dynamics of the treatment effects. This is intrinsically important, as the reorganization of the LHAs is expected to influence population health progressively over time. We specify a baseline regression model that includes municipality and year-fixed effects, and a set of time-varying controls, which identifies the effects under a conditional parallel trends assumption.

Our analysis indicates that, prior to LHA consolidation, mortality patterns followed a similar trajectory across treated and never-treated municipalities. The policy does not exhibit immediate effects. While the policy does not demonstrate immediate effects, significant adverse health outcomes emerge in the medium to long term. Specifically, there is an increase in the total mortality rate of up to 2.4% and an average treatment effect on the treated (ATT) of 1.8%, starting from the fourth year post-implementation. These causal effects are credibly identified

US. Other studies analyzed the difficulties in ensuring political support for consolidation reforms of local public healthcare services by measuring the role of scale economies and differences among local communities (Bates et al., 2011).

[4]For example, Bailey and Goodman-Bacon (2015) measure the health effects of establishing the Community Health Centers, which extend access to primary care to the poor, on mortality rates in US counties. Miller et al. (2021) investigate the relationship between public health insurance eligibility and mortality. Clayton (2019) assesses the effect of Medicaid prescription drug spending on mortality rates at the state level. Mora-García et al. (2024) evaluate the long-term effects of implementing primary healthcare intervention in Costa Rica. Wherry and Meyer (2016) study the long-term mortality effects of Medicaid eligibility during childhood.

as evidenced by a placebo analysis, in which treatment status is randomly reassigned to municipalities by altering the policy year, the regions involved, and the LHAs included within those regions.

The policy's detrimental health effect holds steady across different control group specifications, particularly when the analysis is restricted to treated and never-treated municipalities that are spatially proximate, defined as within a 20 or 15-minute travel distance or being adjacent municipalities. Importantly, our findings remain robust when considering the national government's "Debt Rescheduling Plans" (DRP), a policy implemented during a partially overlapping period. Recent literature has documented remarkable adverse health effects associated with the DRPs (see, e.g., Depalo, 2019; Arcà et al., 2020). Under the DRPs, the central government imposed stringent limits on public health spending, reorganized the hospital system, and introduced co-payments for essential services in regions with substantial health budget deficits. The overall effect on total mortality is only partially attributable to the oldest old population. Significant post-merger variations in mortality are observed within the working-age population (ages 25-49 and 50-64). For individuals aged 0-74, the increase in preventable and treatable mortality nearly doubles that of non-amenable mortality. Furthermore, our analysis indicates that the policy's adverse impact is primarily driven by a substantial rise in mortality related to circulatory system diseases and cancers.

To gain insights into the factors underlying these negative outcomes, we exploit the observed differences in municipality types and explore potential heterogeneity patterns influencing the overall impact. We find that mergers were particularly detrimental for very large LHAs, and more in general in the event of large increases in the population served by these administrative entities following the consolidation process. Efficiency gains, in the form of cost-saving from the centralization of administrative functions, appear to have been achieved at the expense of reducing overall health quality, as reflected by mortality indicators. Further analysis reveals that municipalities whose LHAs changed headquarters after the merger experienced the most significant negative effects. Additionally, by using the Istat classification of inner areas – which differentiates urban and rural municipalities based on access to essential public services – we find that Outlying municipalities, which surround Core urban centers providing all essential services, are the most adversely affected by the policy.

Our analysis relates to the debate on hospital mergers and closures, which has extensively examined both the potential pitfalls and benefits, including productivity gains and cost savings (see, e.g., Dranove and Lindrooth, 2003; Schmitt, 2017), as well as the impact on patients' health outcomes (see, e.g., Buchmueller et al., 2006; Gaynor et al., 2012). Similarly, assessing the effects on health outcomes when mergers involve local health jurisdictions rather than specific healthcare providers is important. Our findings contribute to this discussion by pinpointing the potential unintended consequences of large-scale local health district consolidations, particularly regarding their impact on healthcare effectiveness.

2 Related literature on consolidation in health care services

Consolidation of public health departments is not a common topic in the economic literature, with only a few existing studies mainly focused on estimating the cost-saving effects of these reforms. Efficiency gains arising from consolidation have been found in the case of US local health departments (e.g. Bates and Santerre, 2008; Bates et al., 2011; Hoornbeek et al., 2015; Bernet and Singh, 2015). The estimated minimum efficient scale is usually relatively small. In particular, Santerre (2009) estimated significant efficiency gains from mergers involving up to 100,000 inhabitants. In Italy, where LHAs have primary responsibilities regarding health care provision, the cost-effectiveness impact of the consolidation process has been studied by Di Novi et al. (2018). By focusing on a particular subset of production costs (namely, administrative costs and the costs of purchased goods), the authors found evidence of strong economies of scale, with a minimum efficient scale corresponding to about 1,000,000 LHA enrollees. To the best of our knowledge, only Johar and Savage (2014) have focused on the impact on health care quality, finding overall positive (though heterogeneous) results regarding waiting times arising from the amalgamation of adjacent area health services in New South Wales, Australia.

Although not directly comparable, research on hospital mergers offers valuable insights into the potential health effects of administrative consolidation, presenting mixed evidence. Studies investigating the volume-outcome relationship have provided evidence of enhanced quality in merged hospitals (e.g., Gaynor et al., 2012). This is because a more significant number of treated patients – especially when they share equal or similar health needs – increases the quality of care, according to the "practice makes perfect" hypothesis (Luft et al., 1987; Gaynor et al., 2005).

However, even effective access to care matters, as originally remarked by Buchmueller et al. (2006) and more recently confirmed, among the others, by Avdic (2016), who found adverse effects of closures of neighboring emergency hospitals in Sweden in terms of access to care and increased heart disease mortality. Carroll (2023) has emphasized that hospital closures in rural US areas resulted in lower spending but higher mortality rates, primarily due to increased travel costs, which reduced access to high-quality hospitals. In the same vein, Gujral and Basu (2019) have identified a sharp penalty for rural areas vis-à-vis urban areas. Both beneficial and harmful effects have been found by Avdic et al. (2024) in the case of maternity ward closures on infants' and mothers' health. While closures increased efficiency and improved child outcomes, they adversely impacted maternal health, highlighting challenges with scaling up capacity and managing crowding. At a larger scale, slightly adverse effects in terms of patient experience have been found for a large sample of hospital merger episodes (Beaulieu et al., 2020). Evident adverse outcomes regarding hospitalizations and mortality have arisen in the specific case of consolidation of dialysis facilities within large chains (Eliason et al., 2019).

The evidence on health outcomes from hospital care reorganization provides valuable insights into the potential health outcome variations following the consolidation of local health departments. On the one hand, positive effects from the amalgamation of bordering LHAs may emerge if the merger facilitates better integration of hospital activities across adjacent geographical areas. Additionally, other factors may enhance health outcomes associated with a specific geographical area involved in LHA consolidation. For instance, higher-quality services could result from creating larger healthcare units for outpatient care, thanks to learning-by-doing, selective referral, or quality-enhancing scale economies. On the other hand, potential drawbacks must also be considered. First, the organizational complexity of a single governance structure might exceed manageable limits. Second, the reduction in the number of jurisdictions could hinder the identification and disclosure of healthcare needs, particularly for vulnerable population groups, leaving some needs partially unmet. Lastly, healthcare centralization usually increases the average distance to healthcare facilities, especially for rural residents, while simultaneously reducing access to informal information networks that help individuals navigate and select appropriate care.

3 The Italian LHAs consolidation process

The Italian National Health Service (NHS), established in 1978, is a regionally decentralized system funded by regional corporate taxation, regional top-ups to the national income tax, and the allocation of VAT revenues. This funding structure ensures the provision of a set of legally defined essential levels of care. The primary objective of the NHS is to provide uniform and comprehensive health care nationwide.

The current institutional setting has arisen from a series of reforms initiated in 1992 (Amato - De Lorenzo reform) and implemented through legislative decrees in 1992 and 1993, progressively introducing universal free-patient choice and establishing 21 autonomous Regional Health Services (RHSs). The decentralization of the NHS became fully effective with a constitutional reform approved in 2001, which granted regions greater autonomy in organizing healthcare services, including technological investments, the hospital network, and the diversification and specialization of local supply. This increased autonomy, naturally, came with added responsibilities in terms of controlling healthcare spending and ensuring efficiency in service provision. The national government maintains a general authority over healthcare policy, including defining the essential levels of care and setting overall spending limits.

The RHSs deliver public health care through sub-regional health authorities. In addition to directly managing a large share of all public hospital providers, the LHAs' role extends to broader responsibilities: from planning and managing the delivery of essential healthcare services, including primary care, outpatient clinic care, preventive health programs, social care-related health care for residents in the included municipalities, and public health initiatives, to contracting with general practitioners and pediatricians to manage budget allocation for healthcare services. The origin of the current role of LHAs dates back to 1995, with the creation of 228 Local Health Authorities (LHAs) and the elimination of previous and much more numerous local health Local Health Units (USLs) managed by single (or groups of) municipalities. This change shifted governance from a system where municipal councils attributed management roles, reflecting local political dynamics, to one where LHAs are endowed with legal personality and entrepreneurial autonomy and managed by a general manager appointed by the regional Government.[5] While most regions adopted the reform in 1995, Sardinia and Sicily, as special statute regions, implemented it one year later. The "regionalization" of local health entities was finalized in 1997 when the Veneto region made a minor adjustment to its structure by merging two LHAs. In 1998, the overall number of LHAs was further reduced to 197 as a consequence of the autonomous initiative by the Lombardia region to restructure its RHS into a quasi-market. This reorganization introduced a clear distinction between funding and purchasing entities (the

[5] This change did not necessarily diminish political influence; instead, it reallocated it from the local to the regional level.

LHAs) and service providers (all hospital entities).[6]

In the years after the constitutional reform of 2001, as regions faced financial pressures from population aging and national austerity constraints, several opted to merge smaller Local Health Authorities (LHAs) into larger entities. The principal motivation for this consolidation was the potential for efficiency gains through economies of scale, achieved by broadening the catchment areas for indivisible services and enhancing organizational integration.

Over the years, as illustrated in Figure A1, many regions redefined the distribution of LHAs, resulting in an overall reduction of 96 authorities. A distinctive aspect of this process is that regions acted independently and at different times, as detailed in Table A2 in the Appendix, with the first mergers occurring in 2004 and the most recent ones in 2017, for a total of 12 years of mergers.[7] Figure 1 documents the consequent expansion of the catchment areas of the LHAs involved in the merger process.[8]

Over the past two decades, the Italian NHS has been characterized by another national cost-saving policy that must be considered when investigating the effect of the LHAs consolidation policy on health outcomes. Specifically, the "Debt Rescheduling Plans" (DRPs), enforced by the central government in ten regions, aim to fix financial imbalances and repay debt incurred by regional healthcare budgets. These measures targeted regions with excessive financial deficits, leading to the implementation of specific strategies aimed at achieving both cost containment and health outcome targets, including expenditure control, reductions in healthcare public funding, and changes in the organization and size of the hospital network.[9] The potential impact of the DRPs on public health may be considerable. Recent empirical evidence confirms that regions under a DRP successfully contained costs as expected (Bordignon et al., 2020). However, they experienced adverse health consequences, such as higher mortality, lower hospitalization rates, and increased south-to-north inter-regional patient mobility (Depalo, 2019; Arcà et al., 2020; Cirulli and Marini, 2023; Guccio et al., 2024). The DRPs partially overlapped with the LHA consolidation process. Specifically, six regions – Abruzzo, Calabria, Campania, Lazio, Molise, and Sicily – were placed under the DRPs in 2007, while Puglia followed in 2010. These regions

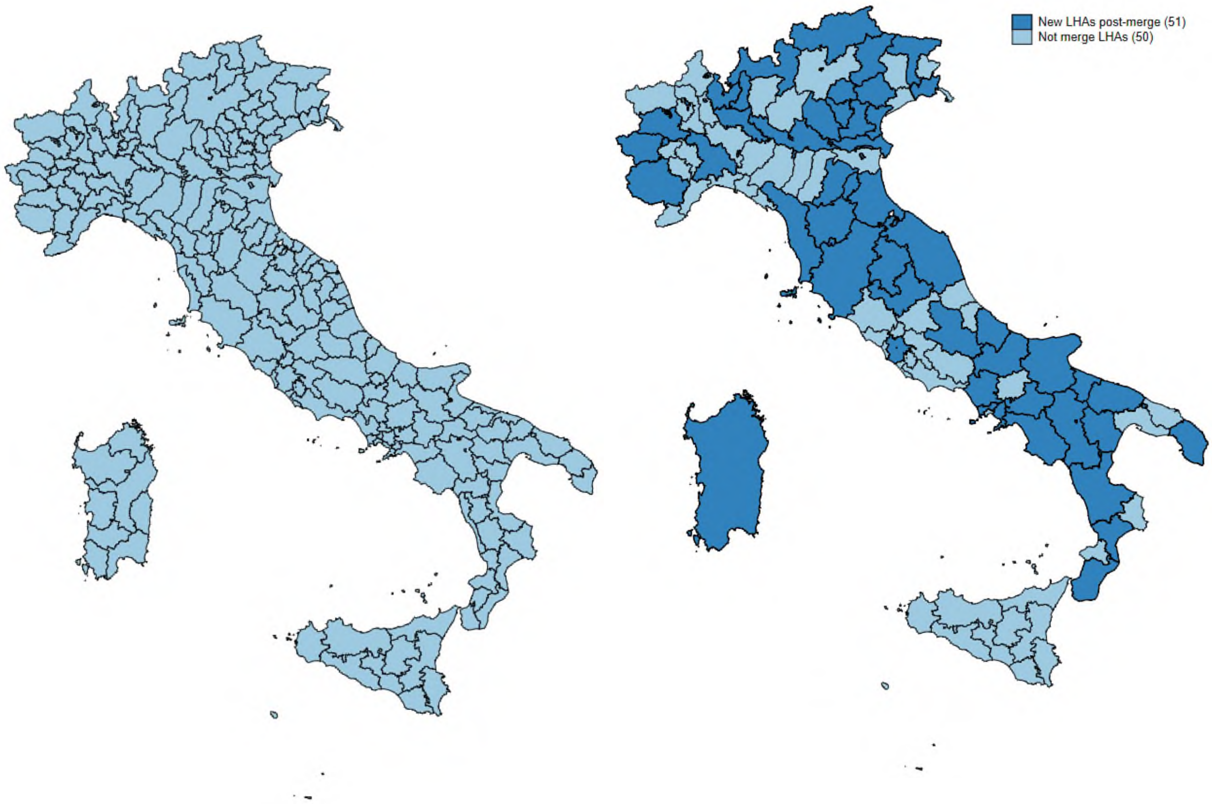
[6] This restructuring not only decreased the number of LHAs in Lombardy from 44 to 15 but also fundamentally transformed the 1995 LHAs by establishing autonomous hospital trusts and Scientific Institutes for Research, Hospitalization, and Healthcare. These changes partially relieved LHAs of direct management responsibilities, refocusing their role on preventive care, health promotion, and healthcare planning.

[7] Only two regions, Emilia-Romagna and Calabria, experienced two waves of mergers involving different LHAs: 2004 and 2014, and 2008 and 2011, respectively.

[8] It is worth noting that, while the transformative reorganization of the RHS in Lombardia in 1998 anticipated the consolidations implemented by other regions after the 2001 reform, the changes from 2004 onward were intended solely to redistribute and expand the LHAs' catchment areas without altering their core mission.

[9] See Arcà et al. (2020) for a thorough description of the DRPs, therein labelled "Recovery plans".

Figure 1: Borders of Italian Local Health Authorities in 1999 and 2018



b. 2018

a. 1999

Notes: Elaborations on Ministry of Health Data. Intra-municipal LHAs borders for Turin, Milan and Naples are not drawn

remained under this regime throughout our analysis period, which extends up to 2018. Liguria and Sardinia were under the DRP regime for three years (2007-2009), while Piedmont was for six years (2010-2015).[10]

4 Data

Our empirical analysis relies on a comprehensive dataset covering Italian municipalities for the period 2002-2018. The Italian NHS transformation described in Section 3 suggests that 2002 is the appropriate year to begin studying the effects of regional consolidation policies introduced by the 2001 constitutional reform, which granted regions greater autonomy in healthcare governance.[11] The effect window we have chosen ends in 2018 to prevent potential confounding effects arising from the COVID-19 pandemic.

Our dataset includes annual inter-census data on resident populations, as well as yearly

[10] As required, in the years following the implementation of the DRPs, targeted regions could re-establish previous health expenditure patterns. Once achieved, the budget equilibrium must be sustained over time.

[11] This choice also ensures at least two pre-policy periods for all municipalities involved in the LHAs' mergers.

municipal-level death counts disaggregated by (5-year) age group, gender, and cause of death (according to the ICD10 classification), as provided by Istat.[12] Each municipality is linked to average taxable income data released by the Italian Ministry of Economics and Finance and is classified according to Istat's definition of inner areas (Core, Outlying, Intermediate, Peripheral, and Very Peripheral municipalities).[13]

To evaluate the impact of the consolidation policy, we associate each municipality with its respective LHA for each year in the period of interest. This process relies on data from the Italian Ministry of Health (MH). Specifically, we identified the municipalities involved in the consolidation using the administrative dataset "Asl–comuni–popolazione" released by the MH in 2023, which details the distribution of citizens enrolled in each LHA by their municipality of residence for the years 1999, 2011, and the period from 2013 to 2018. To account for the years not covered by this dataset, we conducted a documental analysis based on official records released in 2018 by the Ministry of Health and known as "Variazioni ambiti territoriali e ricodifica delle ASL" (variations in catchment areas and LHAs recoding). These records detail the LHA mergers that took place between 2006 and 2017. When this information proved insufficient, we consulted the official NHS hospitals' address book, "Indirizzi delle strutture di ricovero pubbliche e private", released by the MH in 2007. This source allowed us to track the realignment of hospitals that were originally under the same LHA into newly formed LHAs post-consolidation, serving as a reliable indication of municipal involvement in the consolidation process. By systematically monitoring these changes, we were able to map out all modifications in LHAs catchment areas during the period from 2002 to 2018. In particular, for the year 2002, we could confirm that the LHA configuration was consistent with the 1999 official data, as indicated by the 2002 address book that listed the headquarters of 197 LHAs.[14]

As Section 3 outlined, the LHA restructuring process primarily involved merging two or more smaller LHAs into a larger entity. Municipalities within these newly formed LHAs were directly affected by the restructuring and are therefore considered *treated*.[15] A few LHAs experienced

[12]Data on causes of death and age at death are provided separately by Istat, preventing us from analyzing specific age distributions within each cause of death or examining specific ICD-9 causes of death within age groups.

[13]This classification is based on municipalities' accessibility to essential public services (education, health, and mobility), and it is described at https://www.istat.it/it/archivio/273176 and is available at https://asc.istat.it/ASC/.

[14]To validate our methodology, we used the hospital address book to recode the LHAs and verify the correspondence with mergers that occurred between 2006 and 2017, for which we possess official data. Ultimately, we encountered only 10 cases nationwide (all in Apulia) where we could not determine the exact year of change in LHA affiliation. Consequently, these municipalities were excluded from the analysis.

[15]An exception is Lombardy's LHA No. 303, where the catchment area was split into two parts, subsequently merging with two separate LHAs (322 and 323).

jurisdictional changes because of the reassignment of one or more municipalities between neighboring territories. In these cases, we classified the municipalities that changed LHA as treated if they were incorporated into a larger LHA or an LHA of similar size to the original one. Because the reduction in LHA's size could impact local health services provision, we excluded from the analysis all municipalities within LHAs whose population catchment area was reduced by more than 5 per cent due to these reassignments.[16] Additionally, we excluded from the sample any municipalities for which treatment status could not be definitively determined.[17]

4.1 Mortality patterns and summary statistics

To evaluate the medium and long-term health impacts of LHA mergers, we focus on municipal-level mortality rates, a widely recognized indicator of population health. These rates capture the impact of organizational changes across the entire healthcare system, as they reflect not only hospital care quality but also the effectiveness of preventive, primary, and secondary healthcare services managed by LHAs and provided outside hospital settings (see Section 3).

Unlike clinical quality metrics commonly used in research on hospital closures and mergers – which typically focus on mortality rates linked to specific conditions or procedures – population mortality rates provide a more comprehensive indicator of healthcare performance. A limitation of clinical quality metrics is their susceptibility to distortions caused by patient mobility. Specifically, if patients perceive a decline in service quality – such as increased travel distances or reduced availability of specialized inpatient care – they may seek treatment outside their LHA, where they expect higher standards of care. This selective outflow can lead hospitals that attract more complex or severe cases from other LHAs to report worse average clinical outcomes, even if their quality of care remains unchanged. Moreover, hospital-based quality indicators may misleadingly reflect conditions in areas different from those directly affected by LHA consolidation. On the whole, since patients can cross administrative boundaries to seek care, hospital-level data may not accurately represent the healthcare experiences of residents in municipalities exposed to the merger.[18]

[16] This exclusion specifically applies to LHA No. 103, 106, and 107 in Sardinia, LHA No. 311 in Lombardy, LHA No. 104 in Friuli-Venezia Giulia, LHA No. 114 in Veneto, and LHA No. 105 in Calabria.

[17] This includes Rome in Lazio, where mergers involved only part of the city; Turin in Piedmont, which initially consisted of four LHAs, then reduced to two, and finally to one, leading to a distinct treatment regime characterized by a double merger; and fifty-five municipalities within a specific LHA in Calabria (coded as LHA 205), due to a double consolidation within the observed period. Altogether, the exclusion of these cases and the aforementioned reassignments led to the removal of 516 municipalities from our analysis. We note that the results of our regression analysis (available upon request) remain unaffected by including these LHAs.

[18] These issues are particularly relevant in the Italian NHS, where, as in other universal healthcare systems,

Specifically, in our analysis, we calculate the overall mortality rates (henceforth, MR) per 100,000 residents and specific MRs for different groups.[19] The latter include age-specific MRs for the broad categories of 0–14, 15–49, and 50 and older individuals, along with a more granular classification of the adult population: 15–24, 25–49, 50–64, 65–74, and 75 and older. We also calculate cause-specific mortality rates for both the leading causes of death and by distinguishing between amenable and non-amenable deaths. In this case, in line with the classification proposed by Nolte and McKee (2004), we measure amenable mortality by considering a subset of death causes that are potentially preventable through effective public health, primary and secondary prevention measures before disease onset, as well as treatable mortality (referring to deaths that can be avoided through timely and effective healthcare interventions after the onset of a disease).[20]

We begin by examining the evolution of average annual mortality rates over the period of the analysis, distinguishing between never-treated and treated municipalities. Figure 2 displays the annual percentage changes in the overall mortality rate relative to the 2002 baseline, with both time series set to zero in the initial year. Values below zero indicate mortality rates lower than those in 2002, while values above zero signal an increase. This descriptive evidence suggests differences between treated and never-treated municipalities, indicating that further analysis is needed to determine whether these observed trends stem from the consolidation policy. Specifically, Figure 2 reveals an rising trend in total mortality rates, which could partly be attributed to population aging. However, the sharper increase among treated municipalities – particularly from 2007 onward, when the consolidation of LHAs begins in earnest (with a 9% reduction in the total number of LHAs).

patients have the right to seek specialized and hospital care anywhere in the country.

[19] A detailed inspection of the overall MRs reveals several abnormal values. While some anomalies can be attributed to extraordinary events, such as three major earthquakes and a few floods during the period under analysis, others are due to administrative errors in the death registry, particularly in smaller municipalities. To eliminate these outliers, we exclude from the analysis any municipalities that, for at least one year, report an MR exceeding the threshold of mean + $3.5{*}\sigma$. This additionally reduces our sample of 428 municipalities but with an additional loss of population of only 0.37%.

[20] Our measure of avoidable mortality considers Infectious diseases (ICD10 codes: A00-A09 , A37 for ages 0-14; A15-A19, B90, A36, A35, A40-A41, A80 for ages 0-74); Neoplasms (cancers) (ICD10 codes: C18-C21, C44, C50, C53, C62, C81 for ages 0-74; C54-C55, C91-C95 for ages 0-44); Endocrine, nutritional and metabolic diseases (ICD10 codes: E00-E07 for ages 0-74; E10-E14 for ages 0-49); Diseases of the nervous system (ICD10 codes: G40-G41 for ages 0-74); Diseases of the circulatory system (ICD10 codes: I05-I09, I10-I13, I15, I60-I69 for ages 0-74; 50% of deaths for I20-I25 for ages 0-74); Diseases of the respiratory system (ICD10 codes: J00-J09 for ages 1-14; J20-J99, J10-J11, J12-J18 for ages 0-74); Diseases of the digestive system (ICD10 codes: K25-K27, K35-K38, K40-K46 for ages 0-74); Diseases of the genitor-urinary system (ICD10 codes: N00-N07, N17-N19, N25-N27, N40 for ages 0-74); Perinatal mortality (ICD10 codes: O00-O99, P00-P96, A33, A34 for all ages; Q20-Q28 for ages 0-74); External causes (ICD10 codes: Y60-Y69, Y83-Y84 for all ages).

Figure 2: Annual percentage changes in total mortality rate relative to 2002 by treatment status.

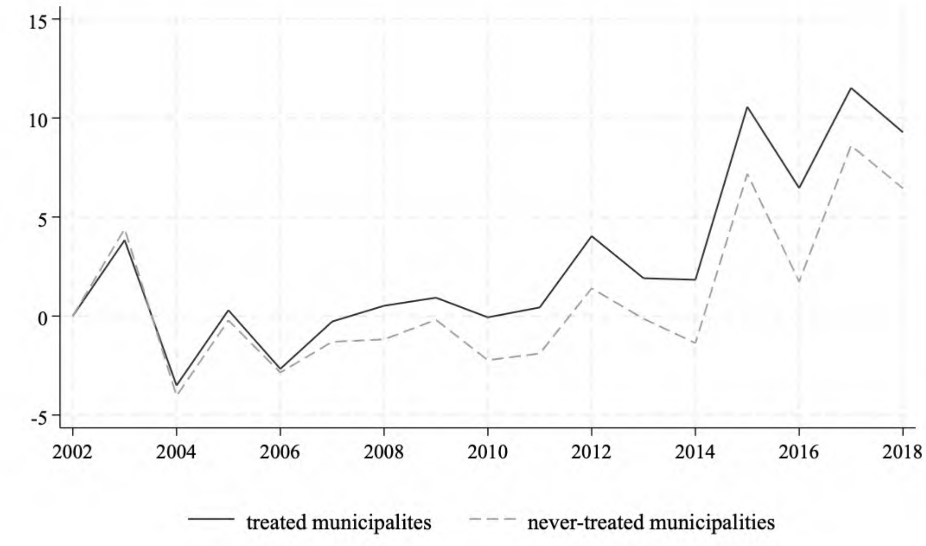
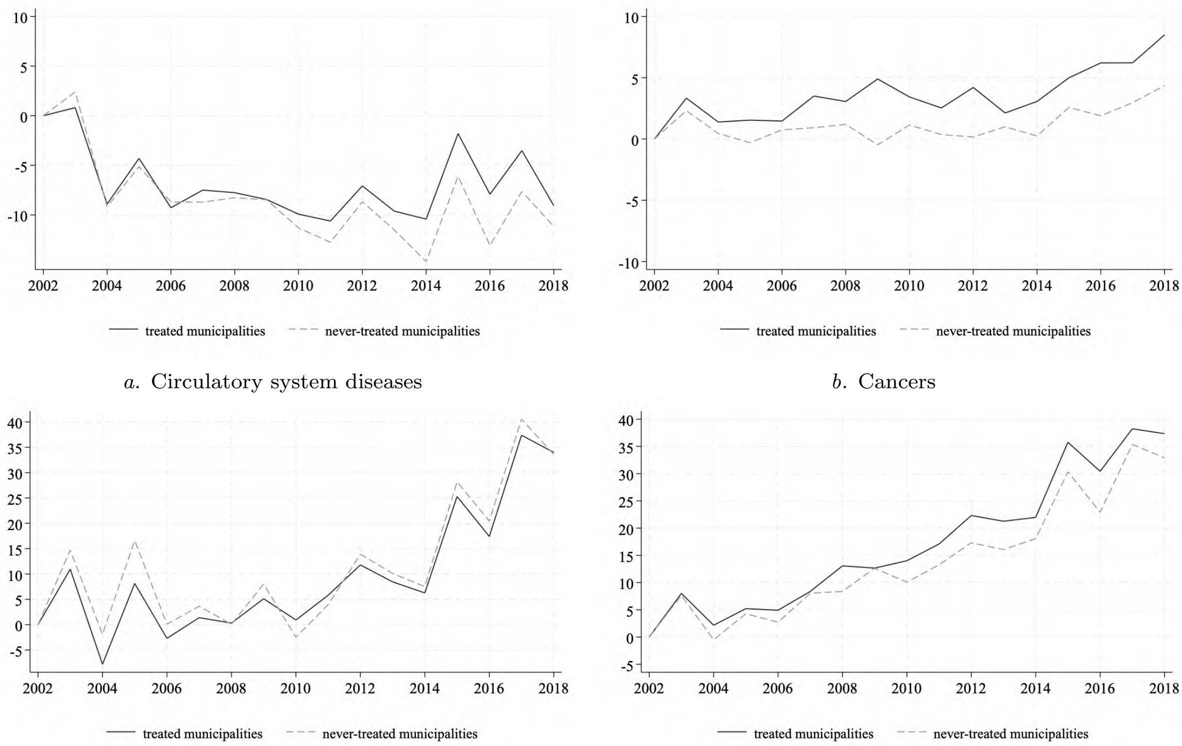


Figure 3: Annual percentage changes in mortality rates for causes of death relative to 2002 by treatment status.



a. Circulatory system diseases

b. Cancers

c. Respiratory system diseases

d. Other causes

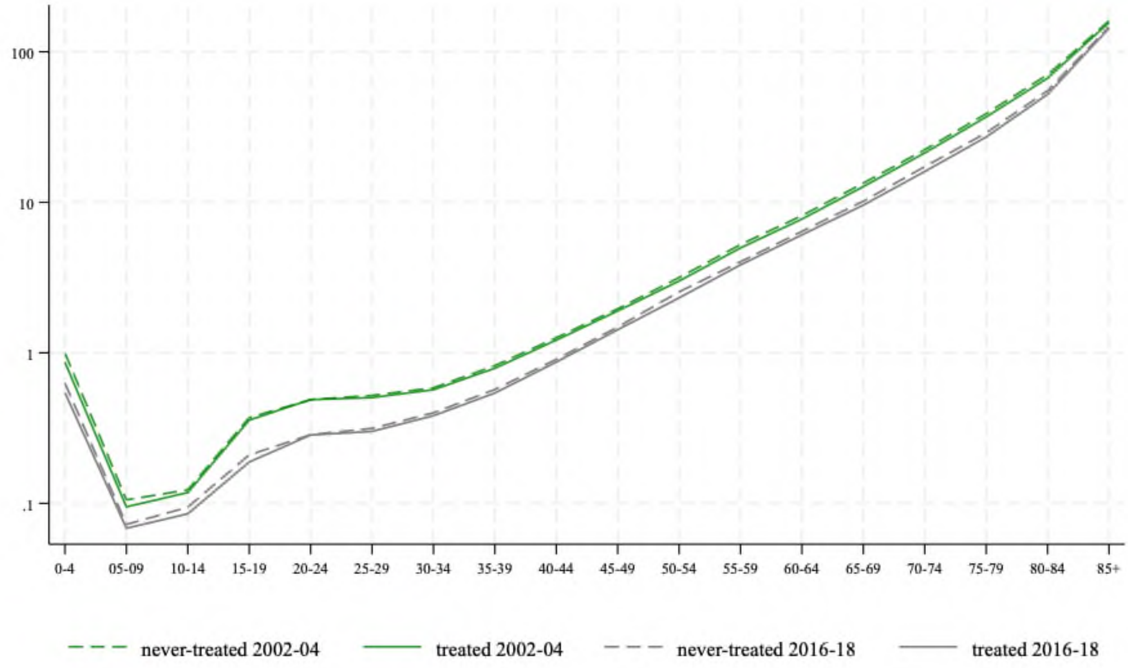
Figure 3 breaks down the relative annual changes in mortality by leading causes of death. Circulatory diseases mortality remains consistently below the 2002 baseline, yet treated municipalities show a steeper increase from 2007 indicating a relative worsening compared to never-treated municipalities. For Cancers, only treated municipalities exhibit a gradual rise in mortality. Respiratory disease and other-cause mortality exceed the baseline from 2008 onward, regardless of policy exposure. While respiratory disease mortality converges between treated and never-treated municipalities, mortality from all other causes rises more markedly in treated municipalities from 2007 onward.

Figure 4 presents changes in age-specific mortality over the analysis period by comparing average mortality rates in the initial (2002–2004) and final years (2016–2018), disaggregated by treatment status. The patterns in the figure follow a J-curve or "hockey stick" shape, which is typical for age-specific mortality rates. The plot highlights a general decline in mortality over time, particularly among younger age groups, for both treated and never-treated municipalities, likely reflecting broader improvements in public health, healthcare access, and living conditions. However, the reduction is less pronounced in older age groups, where mortality rates remain high. Notably, the trends for treated and never-treated municipalities appear largely parallel, suggesting that, at least in aggregate, LHAs consolidation did not lead to substantial divergence in age-specific mortality patterns. Further analysis is required to account for the staggered adoption of the policy and to assess whether heterogeneous effects exist across subpopulations.

To conclude this section, we compare summary statistics for mortality outcomes and the main observable characteristics used in the analysis for treated municipalities, grouped by year of treatment (i.e., by cohort), and for all treated and all never-treated municipalities (Table A1). Dragged by population aging, all cohorts involved in the amalgamation process show a higher mortality rate in the post-policy period compared to the pre-policy period. However, the last two columns of the table show higher mortality rates in municipalities that were never treated compared to those that received treatment over the entire period from 2002 to 2018. While this difference may be partly attributed to a slightly higher proportion of older individuals in the case of the overall mortality rate, it arises even when looking at age-specific mortality rates. Overall, Table A1 supports the use of transformed mortality rates in the econometric analysis, as this adjustment accounts for initial differences in mortality levels at the start of the observation period and helps mitigate potential convergence effects.[21] The table also shows that

[21]Improving mortality rates may be easier when starting from a higher baseline than from a lower one.

Figure 4: Average age-specific mortality rates in 2002–04 and 2016–18 by treatment status.



Notes. The figure shows mortality rates per 1,000 inhabitants, using 5-year age groups from 0-4 to 85 and above. Data are plotted on a logarithmic scale.

cohorts differ in per capita taxable income and age composition of the population. While never treated and treated and are very similar with respect to per-capita income on average, large differences emerge at the cohort level. The same applies when considering the age structure. This motivates using per-capita income and population shares as time-varying covariates to guarantee the conditional parallel trend requirement in the econometric analysis.

5 Empirical strategy

We leverage the variation in the LHA consolidation process across and within regions to identify the causal impact on mortality by means of a Difference-in-Differences (DiD) approach. While some regions implemented the consolidation policy, others did not. Among the regions that adopted the policy, the timing of amalgamations varied. Moreover, not all LHAs within adopting regions were involved in the mergers, resulting in differential exposure to the policy at the municipal level. This staggered rollout quasi-experimental design allows us to group municipalities by the year of treatment, creating twelve distinct treatment cohorts over the period from 2002 to 2018. Municipalities affected by the LHA consolidation during this interval remained under treatment until the end of the observation window.

Due to the staggered adoption of the policy, treated municipalities exhibit variation in treatment exposure duration. This implies that the estimation of the overall static average treatment effect on the treated (ATT) is affected by differences in treatment variance across cohorts rather than solely by the number of units adopting the policy at the same time. For this reason, we implement an event-study design that compares treated and never-treated municipalities. By aligning observations at specific leads and lags, each cohort's timing is standardized relative to their treatment initiation, thus effectively controlling for differences in treatment exposure duration. Moreover, analyzing the full dynamics of the treatment effect on mortality rates is intrinsically important, as the reorganization of the LHAs is expected to influence population health progressively over time. The event-study approach allows us to analyze the policy's impact in detail, capturing both its immediate effects and the progression of its influence on mortality rates over the medium-to-long term.

We estimate the following linear regression model:

$$y_{it} = \alpha_i + \lambda_t + \sum_{l=-K}^{L} \beta_l D_{it}^l + \delta X_{it} + \epsilon_{it} \tag{1}$$

where y_{it} represents the mortality rate per 100,000 residents for municipality i in year t, as described in Section 4. The terms D_{it}^l are binary indicators denoting whether municipality i is l time periods before or after initial treatment ($l = 0$ denotes the treatment period), whereas the coefficients β_l capture the leads (pre-treatment) and lag (post-treatment) effects. The fixed effects α_i and λ_t account for municipality-specific and time-specific heterogeneity, respectively.

In the baseline specification, all never-treated municipalities serve as control group. To assess the robustness of the results under different definitions of the control groups, we restrict the sample based on geographical proximity. Specifically, we select treated and never-treated municipalities that are geographically adjacent or within a 15-minute or 20-minute travel distance. Furthermore, we redefine the control group by using only last-treated municipalities as controls.[22]

The traditional approach for estimating equation (1) is the two-way fixed effects estimator (TWFE). In the event of heterogeneous treatment effects in time for the cohorts, however, the TWFE may yield uninterpretable and misleading estimates of the causal effects, as a consequence

[22] Consolidation at the LHA level precludes policymakers from selectively involving a single municipality in the merger process. Therefore, a propensity score matching approach to select control municipalities is not appropriate.

of non-convex and non-zero implicit weighting, and it invalidates pre-trend tests (Goodman-Bacon, 2021; Sun and Abraham, 2021). To address these issues, a large set of alternative estimation approaches has been developed in recent years.[23] In this analysis, we mainly employ the approach proposed by Sun and Abraham (2021), known as the Interaction Weighted (IW) estimator, which identifies cohort-time ATTs and averages them across cohorts at a given l. Specifically, the IW method proceeds in two stages. In the first stage, the ATTs for each cohort ($CATT$) are estimated using a regression model where cohort and relative time indicators are interacted. In the second stage, the event-study coefficients are calculated as a weighted average of these $CATT$s, using cohort shares as weights. These weights explicitly account for the number of units adopting the policy at the same time.[24]

In our application, the interacted regression in the first stage is estimated via Ordinary Least Squares (OLS) using the total municipal-level population in 2002 as weight, thereby reflecting differences in the relative impact of the policy. To account for heterogeneity in baseline outcome levels across municipalities and address skewness in the distribution of mortality rates, we generally use the inverse hyperbolic sine transformation. Under this specification, the estimated β_l are interpreted as proportional changes in mortality.

The IW estimation procedure incorporates time-varying covariates into the regression model. Our model accounts for several municipal-level covariates X_{it}, including population shares of different age groups, average income, and an indicator for regions where local health services were exposed to DRPs.[25] We include the shares for the age groups 50–64, 65–74, 75–84, and 85 and above to account for changes in the composition of the population over time.[26] Controlling for the evolution of age composition is crucial when employing crude mortality rates, which is necessary when age-adjusted mortality rates are unavailable[27] and when control variables included in the model cannot be adjusted for the same population compositional effects (Bauernschuster et al.,

[23]We can broadly distinguish between approaches based on weighting up cohort-specific ATT (e.g., Sun and Abraham, 2021; Callaway and Sant'Anna, 2021), focusing on 'first-time switchers' (de Chaisemartin and d'Haultfoeuille, 2020), using "imputation" estimators of counterfactual outcomes for treated units (Gardner, 2022; Borusyak et al., 2024; Wooldridge, 2021), and stacking data through balancing in relative event time (Cengiz et al., 2019).

[24]The larger the number of units simultaneously exposed to the policy, the greater their influence on the aggregate estimates of the treatment effects.

[25]Without covariates, using the IW estimator is equivalent to adopting the Callaway and Sant'Anna (2021) approach with never-treated units as controls. In fact, the latter approach assumes parallel trends conditional only on time-invariant covariates, and relies on common support when the generalized propensity score is used.

[26]In the excluded age group (below the age of 50), deaths account for only about 3.8% of total mortality in the period 2002-2018. In contrast, deaths of individuals aged 50–64 constitute approximately 8.8%, those of individuals aged 65–74 account for 15.2%, those of individuals aged 75–84 represent 32.2%, and those of the oldest old (85 years old and above) are approximately 40% of the total observed mortality.

[27]As noted in Footnote 12, age-adjusted rates cannot be calculated for cause-specific or amenable mortality rates, given the absence of detailed age-at-death information in the provided data.

2019; Rosenbaum and Rubin, 1984). By controlling for average taxable income, we aim to capture differences in socioeconomic status that may influence health outcomes before and after the consolidation policy. Higher-income populations have lower health risks (e.g., through better education and healthier behaviors) and easier access to private healthcare services. Wealthier municipalities might also have more and better public services and infrastructure that ensure easier and quicker access to healthcare. Finally, we have included the DRP indicator to account for differences in mortality trends between municipalities in regions that were forced to contain their health costs by focusing on specific targets such as the number of hospital beds and the organization of hospital services (Depalo, 2019).

Contingent upon the controls mentioned above that are expected to affect mortality rates, we are confident that the conditional parallel trend assumption holds. In fact, as shown by Figure A2, the distributions of these covariates overlap for treated and never-treated units, thus ensuring that municipalities are exposed to treatment independently of those covariates and that our never-treated municipalities are a valid control group for the treated group. In particular, in 2002-2018, about 39.7% of never-treated municipalities also experienced the DRP, compared to about 27% among treated municipalities. To further examine the potential role of DRPs, whose implementation partially overlaps with the amalgamation of LHAs throughout the country, we conduct separate estimates for municipalities with and without a DRP. This approach tests whether mortality drifts appear even in areas unaffected by the related additional cut in health care expenditures.

In the next section, we estimate equation (1) for the total and age-specific mortality rates. Additionally, we present a simple decomposition of the policy impact by focusing on causes of death. When analyzing age groups or specific causes of death at the municipal level, a trade-off arises between modeling proportional changes while accounting for cohort heterogeneity and the presence of mass at zero, which makes the use of inverse hyperbolic sine (or logarithmic) transformations incorrect (Chen and Roth, 2024). For instance, in 2018, 86% of municipalities recorded zero deaths in the age group 0-14. In such cases, we settle for the Poisson pseudo-maximum likelihood (PPML-) TWFE estimator with mortality rates in levels. While the estimated coefficients can still be interpreted as percentage changes in the mortality rate, this approach shares with the linear TWFE the limitation of not accounting for treatment effect heterogeneity across cohorts and over time.[28]

[28]Moreover, municipalities with zero deaths for all years are excluded from the analysis, significantly reducing the sample size (this mainly occurs when considering mortality among the 0-14 and 15-24 age groups). This

6 Results

6.1 Effects on total mortality

This section presents yearly event-study estimates of equation (1) and discusses the dynamic effect of the LHA mergers on population mortality rates at the municipal level. The baseline specification includes municipality and year-fixed effects and time-varying controls and relies on the assumption of conditional parallel trends. This means that the evolution of the mortality outcome is the same for treated and untreated units conditional on population shares for specific age groups, per capita income, and being located in a region under DRP. The model specification considers an effect window between $l = -6$ and $l = 9$, which provides a left endpoint sufficiently far from the time of consolidation to verify the absence of pre-trends and an upper limit that avoids considering very long-run effects estimated only with a small number of treated municipalities).[29]

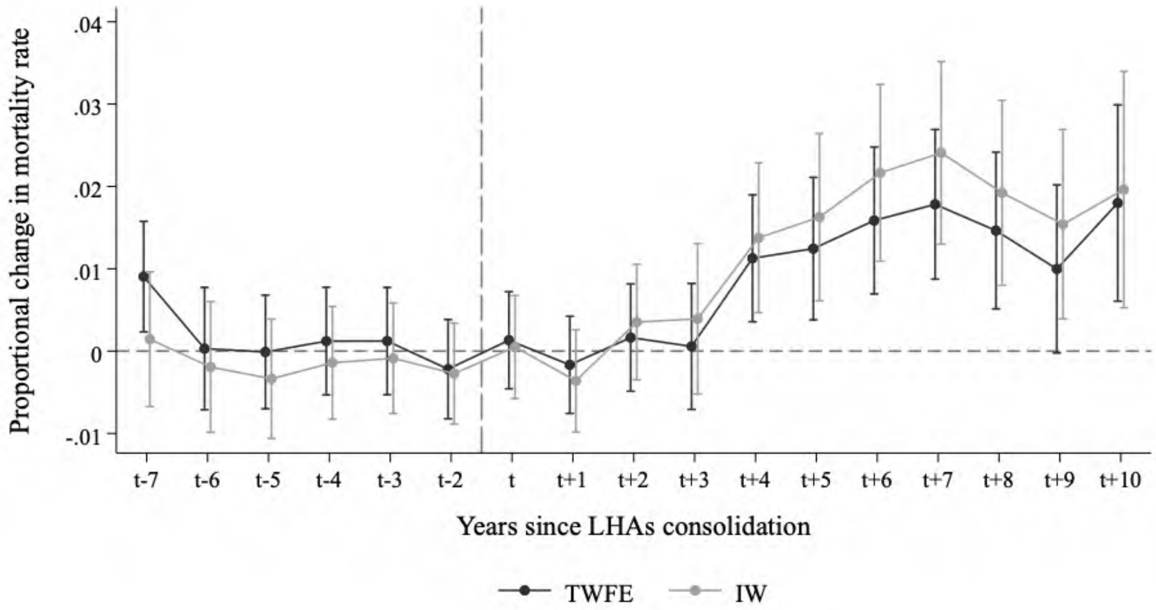
We begin examining the population-wide health effects of the LHA consolidation process by focusing on the total mortality rate (for all ages and all causes). Figure 5 compares estimates obtained using the TWFE and the IW estimator. As discussed in the previous section, the latter estimator addresses the issue of treatment effects in specific periods being contaminated by effects from other periods in the event of group heterogeneity. This ensures that estimates of treatment leads can serve as a proper test for parallel pretrends. As we can see, conditional on the time-varying covariates described in Section 5, we do not observe any differential trend in the years preceding the LHA mergers. While both lines in the plot indicate a positive and statistically significant mortality drift starting four years after the merger, the TWFE estimates are smaller, suggesting an underestimation of average treatment effects at specific bins since the consolidation occurred. According to the IW estimation results, relative to the average of 993.2 deaths per 100,000 inhabitants in the year before the reform, the percentage increase in the mortality rate in the post-treatment period ranges between 1.4% (at $t + 4$) and 2.4% (at $t + 7$). As reported in Table A3, the ATT in the entire post-treatment period is a proportional increase in the death rate of 1.15%. This rises to 1.84% when considering the restricted effect window $[t + 4, t + 9]$ during which the policy has a significant and persistent impact on mortality.[30]

exclusion is necessary because, in the presence of statistical separation, the maximum likelihood estimates for Poisson models do not exist.

[29] At this purpose, we bin the data at $l = -7$ and $l = 10$, instead of estimating a fully dynamic specification. The estimates in the effect window [-6, 9] are unaffected by this choice, as shown in Figure A3.

[30] As an additional check, Table A3 also reports the estimates obtained by means of the Gardner (2022) "two-

Figure 5: Effects of LHAs consolidation on total mortality rate

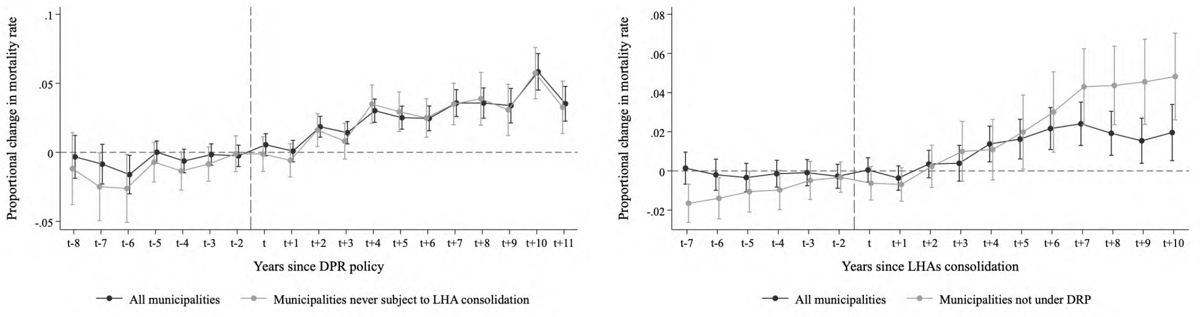


Notes: Dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018 (see Equation 1) for the TWFE estimator and the IW estimator by Sun and Abraham (2021). The TWFE and interacted regression are estimated by OLS using population in 2002 as weights, including municipality-fixed and year-fixed effects. The relative time before the consolidation $l = -1$ is excluded and indicated by the vertical line. The left endpoint is binned at $l = -7$, and the right endpoint is binned at $l = 10$. The dependent variable is the inverse hyperbolic sine of the mortality rate per 100,000 residents calculated for all ages and causes of death. The models include time-varying controls: the shares of individuals aged 50-64, 65-74, 75-84 and 85 and older, per capita income, and a binary indicator of whether municipalities are in regions under a Debt Rescheduling Plan. Standard errors are clustered at the municipality level. Confidence intervals are indicated at the 95 per cent level.

We address the potential confounding effect of the DRPs, a policy that partially overlapped the LHAs consolidation and has well-documented effects on mortality at the regional level (Depalo, 2019; Arcà et al., 2020; Cirulli and Marini, 2023; Guccio et al., 2024). Our analysis confirms the detrimental effect of DRPs on municipal-level mortality rates and, crucially, demonstrates that the growing trend in mortality rates cannot be solely attributed to the concurrent implementation of DRPs, as the evidence reveals that LHAs mergers independently contribute to the observed increase in mortality. Panel a of Figure 6 presents staggered event-study estimates from a model specified in equation 1, with $l = 0$ corresponding to the onset of DRPs. The plot compares the sample that includes all municipalities and the subsample of those exposed to LHAs consolidation at some point, which could potentially confound the causal effect of DRPs. The results provide compelling evidence, in an event-study design, of the harmful effects of DRPs on public health. For municipalities involved in DRPs, mortality rates increase by approximately 2.7% on average after implementing these plans. Panel b shifts focus to the distinct

stage DiD" estimator, which is well-suited for considering time-varying controls. The results, as shown in the third column of the table, are mostly confirmed, with a slightly smaller ATT in the $[t + 4, t + 9]$ effect window.

Figure 6: Effects of LHAs Consolidation and Debt Rescheduling Plan policies on the total mortality rate



a. DRP policy *b.* LHAs consolidation

Notes: In each figure, dots indicate the estimated coefficients for each year before and after the reform: Debt Rescheduling Plan (Panel *a*) and LHAs consolidation (Panel *b*) using data for the period 2002-2018 (see Equation 1). The IW estimator by Sun and Abraham (2021) is used. All models include time-varying controls: the shares of individuals aged 50-64, 65-74, 75-84 and 85 and older, and per capita income. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

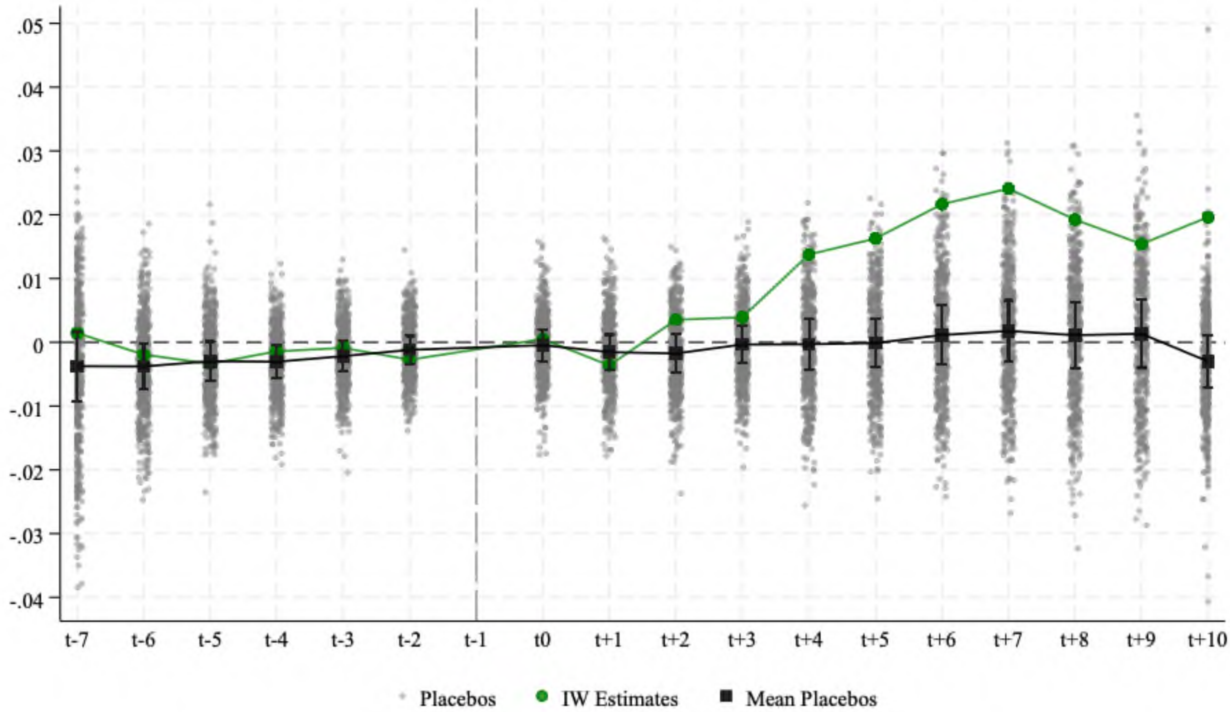
impact of LHAs consolidation. The baseline results, already reported in Figure 5, are compared with those derived from the subset of municipalities that were never subject to a DRP. The findings underscore the independent and substantial role of LHAs consolidation in driving mortality outcomes. For municipalities unaffected by DRPs, the mortality trend steepens markedly from $t+5$, suggesting that the adverse impact of LHAs amalgamations is confirmed and amplified in the absence of concurrent DRPs. [31]

The direction and intensity of the consolidation's impact on mortality are somewhat unexpected. In view of that, we are particularly concerned that the rejection of the null hypothesis of no effects may be driven by unobserved trends – such as population aging not adequately accounted for through population shares – that are not captured by time-fixed effects or unobserved heterogeneity. To address these concerns and assess the robustness of our findings, we conduct a placebo analysis by simulating alternative patterns to the observed policy. Specifically, we generate 500 "fake merger policies" that replicate the nature of the actual policy while maintaining its essential structure: the policies affected only groups of municipalities (LHAs) and never individual municipalities. Treatment status is randomly assigned to municipalities by varying the policy year, the regions involved, and the LHAs within those regions. This ensures that all municipalities within an LHA are either treated - if their LHA is merged with another - or untreated, otherwise.[32] We estimate our main model for each of the 500 fake policies and

[31]Table A3 reports event-study coefficients for the model estimated on the subsamples of municipalities in regions under the DRP and those never under the DRP.

[32]Given that fake treatment times and fake treatment units are both considered, our exercise represents a "mixed placebo" test.

Figure 7: Placebo estimates of LHAs consolidation on total mortality rate



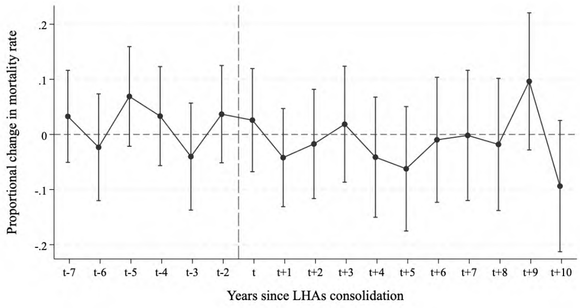
Notes: Small dots represent point estimates for 500 "fake merger policies" randomly generated. The black connected squares indicate the mean of the placebo estimates; the red connected circles represent the point estimates obtained for the actual consolidation policy. The IW estimator by Sun and Abraham (2021) is used for all models. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

examine whether these simulated mergers detect a statistically significant change in mortality rate. Figure 7 illustrates the results: the effects of the actual consolidation policy (the red connected circles), the point placebo estimates (the small grey dots), and the average of the placebo estimates (the black connected squares). As the figure shows, the fake amalgamations neither significantly affect mortality nor exhibit significant pre-trends. This confirms the specificity of the effects estimated for the actual amalgamation policy.

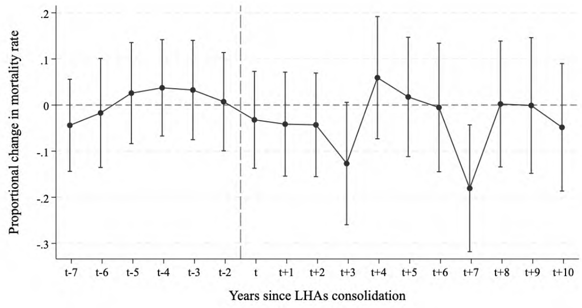
6.2 Effects on age-specific mortality

We further investigate the effects of the consolidation policy across specific age groups. Figure 8 and Table A4 present event-study estimated coefficients for age-specific mortality rates, calling attention to key demographic segments: the youngest population (ages 0–14), the working-age population (subdivided into young adults aged 15–24, adults aged 25-49 and older adults aged 50–64), the retirement-age population (ages 65-74), and the oldest old population (ages 75 and above). This granular categorization reveals notable heterogeneity in both the magnitude and timing of the policy's effects across age groups. Given mass at zero for age-specific mortal-

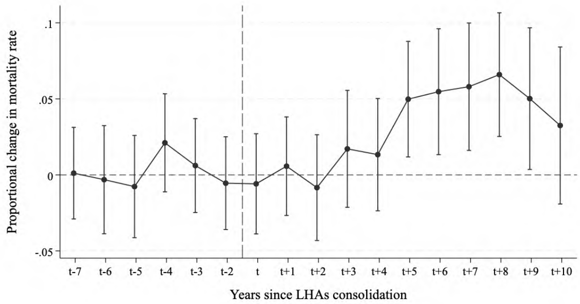
Figure 8: Effects of LHAs consolidation on age-specific mortality rates



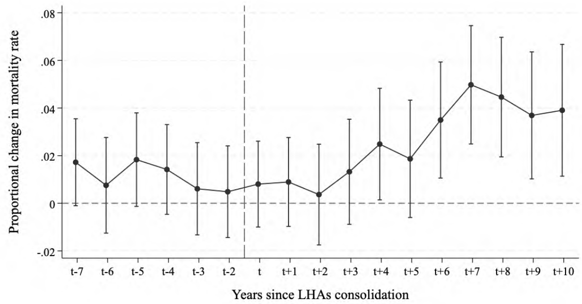
a. age 0-14



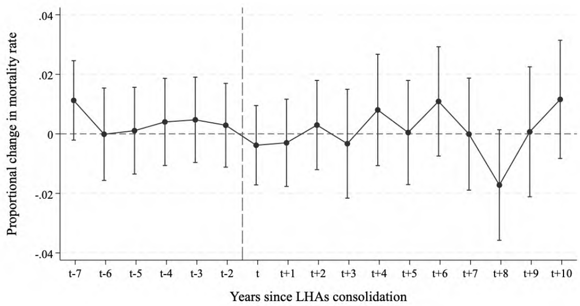
b. age 15-24



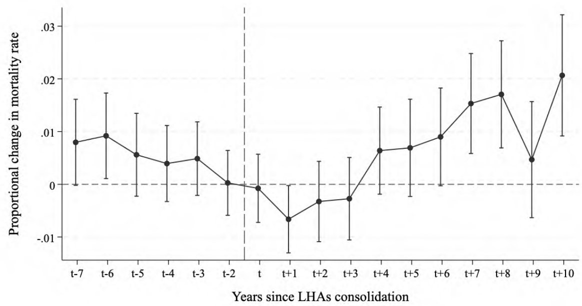
c. age 25-49



d. age 50-64



e. age 65-74



f. age 75 and older

Notes: In each figure, dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018 (see Equation (1). The PPML-TWFE estimator is used for all models, which include total population in 2002 as weight, municipality-fixed effects and year-fixed effects. The dependent variables are age-specific mortality rates per 100,000 residents. All models include time-varying controls: per capita income and a dummy variable indicating whether the municipality is in a region under a Debt Rescheduling Plan. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

ity at the municipal level, we retain the age-specific mortality rates in levels and apply the PPML—TWFE estimator as discussed in Section 5.

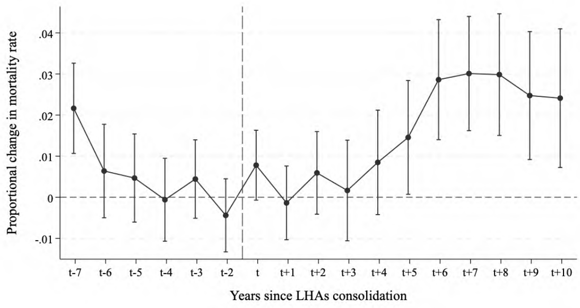
In particular, we find that the general effect on total mortality is mainly driven by the working-age population (ages 25-49 and 50-64), and partially by the oldest old population. For the youngest sub-populations (ages 0-14 and 15–24) and adults of retirement age (65-74), the estimates suggest that, on average, the effect on the treated is null. Specifically, the age group 25–49, as reported in Panel *c*, is characterized by more stable and sustained increases in mortality rates after the mergers, which begin at $t + 5$ and continue through $t + 8$ with a peak of about 6.6% relative to a baseline of 83 deaths per 100,000 inhabitants of the same age category). For older working-age adults (50–64 years old), Panel *d* shows an earlier effect of the policy at $t+4$, but the effect is larger from $t + 6$ onward, reaching a peak of 5% with respect to the baseline rate of 431 deaths per 100,000 inhabitants of the same age. The detrimental effects of mergers are also evident among the oldest segment of the population (75 and older). Panel *f* displays significant coefficients seven and eight years since the policy, with an increase of about 1.5-1.7%. However, this percentage must be contextualized by considering the substantial contribution of this age group to the total number of deaths. Specifically, the baseline mortality rate for individuals aged 75 and older stands at 7,110 deaths per 100,000 inhabitants, underscoring the amplified impact of this age group on overall mortality figures.

With the data currently available, we cannot provide a definitive explanation of the mechanisms involved. Potential mechanisms for these effects may include the disruption of healthcare services, financial pressures arising from labor market fluctuations, and the specific health and healthcare needs of these age groups, which seem to be inadequately addressed following the LHA consolidation.

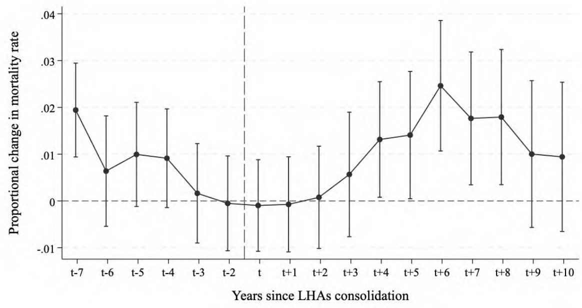
6.3 Effects decomposition by causes of death

To unfold the underlying drivers of the policy's impact on total mortality, we estimate equation (1) using mortality rates calculated for specific types of deaths. The total mortality rate is disaggregated into major causes of death and amenable and non-amenable causes, as defined in Section 4. Estimations are again based on PPML-TWFE, due to zero deaths at the municipal level. The analysis in Figure 9 and Table A5 highlights that the adverse impact on total mortality is primarily driven by increased deaths caused by circulatory system diseases and cancers. This may point to disrupted critical healthcare services required for managing acute and chronic cardiovascular conditions, for which primary care providers play a crucial role in early detection and management and for which hospital care is essential for delivering specialized and timely treatments. Specifically, for deaths related to circulatory system diseases (Panel *a*), the effect is stable and consistent over time, starting from $t + 5$ and peaking in the seventh year since the policy when the mortality rate increases by 3%. Panel *b* shows an earlier positive effect on deaths due to cancer, at $t+4$, lasting until $t+8$, with the largest increase (+2.5%) six years since the reform). Results for mortality related to digestive system diseases (Panel c) are less clear, as effects statistically different from zero emerge only at four and six years since the policy, with coefficients of the opposite sign and close size.

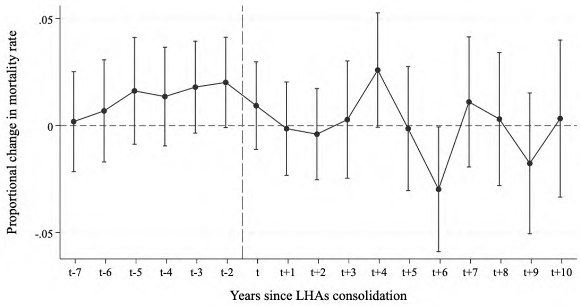
Figure 9: Effects of LHAs consolidation on mortality rate by causes



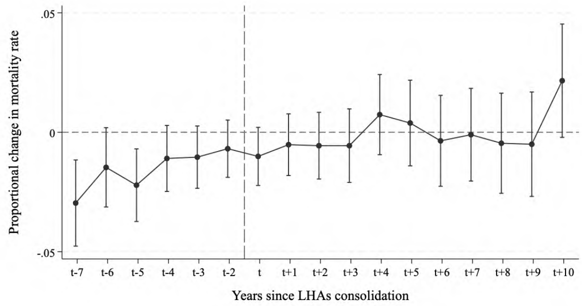
a. Circulatory diseases



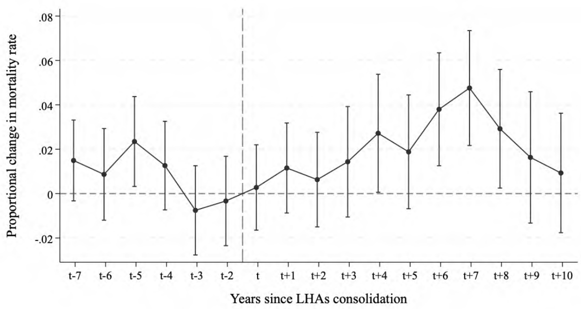
b. Cancer



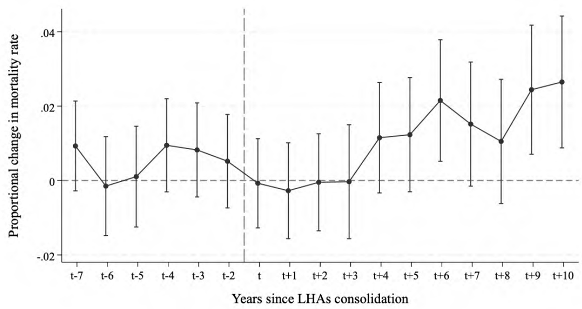
c. Digestive diseases



d. All other causes



e. Amenable causes



f. Non amenable causes

Notes: In each figure, dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018 (see Equation (1)). The PPML-TWFE estimator is used for all models, which include total population in 2002 as weight, municipality-fixed effects and year-fixed effects are included. The dependent variables are mortality rates per 100,000 residents calculated for the causes of death that explain more than 10% of total deaths, and for amenable and non-amenable causes of death in the population aged 0-74. The amenable and non-amenable mortality models do not include the shares of individuals aged 75-84 and 85 and older. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

Effective health policies should aim to reduce the number of deaths related to preventable and treatable conditions. For this reason, using the amenable mortality rate as an outcome measure is particularly relevant. Panels e and f of Figure 9 present the estimated effects on amenable and non-amenable mortality in the population aged 0-74.[33] The detrimental effect of the mergers is clear for both types of mortality and is particularly concerning for preventable and treatable deaths. The mortality rate in Panel e proportionally significantly increases six to eight years since the reform, with a pick of approximately 4.6% and an overall ATT in the effect window $(t+4; t+9)$ of 2.8% (see Table A5). Non-amenable mortality shows a smaller increase of about 2% six and nine years since the reform, with an ATT in the same window of 1.5%.

6.4 Alternative control groups

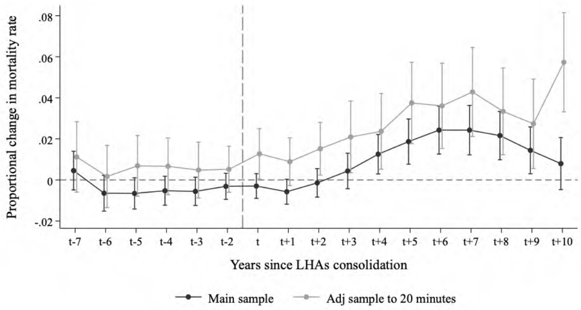
We assess the robustness of our results by enhancing the degree of similarity between treated and control units in several ways. Specifically, we restrict the sample to include only treated and never-treated municipalities that are spatially proximate, using three distinct definitions of proximity: municipalities within a 20-minute travel distance (Panel a), municipalities within a 15-minute travel distance (Panel b), adjacent municipalities (Panel c). These proximate treated and never-treated municipalities are likely to be highly comparable in terms of geographical characteristics and socioeconomic factors that influence mortality patterns. Neighboring territorial units tend to share common features, such as physical geography, weather conditions, air pollution, demographics, and other factors that may affect mortality rates and trends. Panels a to c of Figure 10 present event-study estimates for the total mortality model, comparing the main sample to the three sub-samples defined above. The findings confirm the policy's detrimental effects on mortality, with the effects being sharply large, particularly for the sample of municipalities within a 20-minute travel distance.

As a further robustness exercise, we use the municipalities treated in 2017 as the control group (rather than as treated units), as this cohort represents the last to be treated. Our conjecture is that all treated municipalities share common characteristics that have increased their likelihood of treatment.[34] This approach entails restricting the sample by excluding never-treated units and omitting the years for which $CATT$ estimates are unavailable (2017 and 2018). Panel d of Figure 10 compares event-study estimates for the total mortality model using never-treated

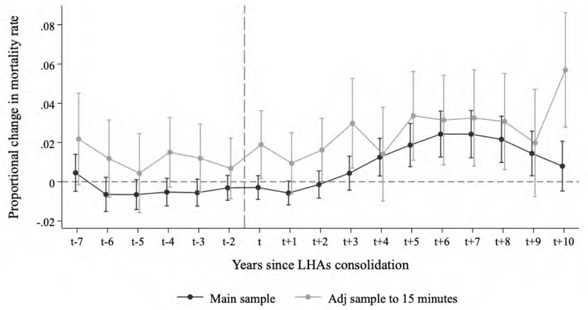
[33]Non-amenable causes account for nearly all deaths (93%) in the total population but represent only 21% of deaths in the 0-74 age group, allowing a more precise focus on the impact of healthcare interventions for this specific segment of the population.

[34]For example, they belong to LHAs exposed to high cost-containment pressures, which led to consolidation.

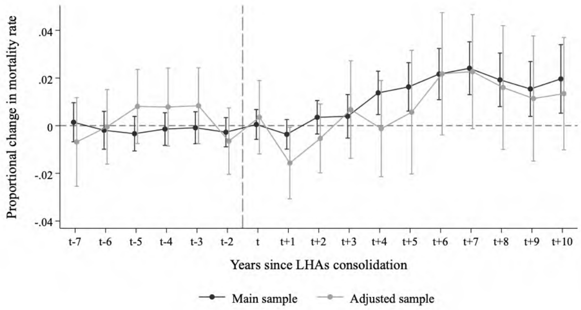
Figure 10: Effects of LHAs consolidation on total mortality rate using sample restrictions



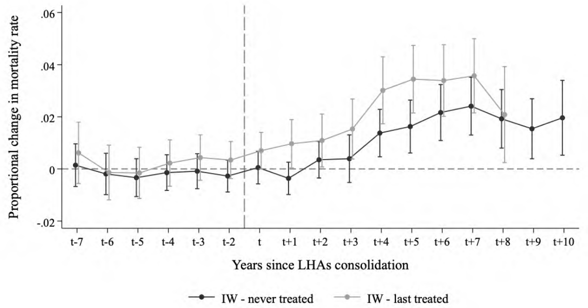
a. Municipalities within 20 minutes



b. Municipalities within 15 minutes



c. Adjacent municipalities



d. Only treated municipalities

Notes: Dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018, except for Panel *d* that uses data for the period 2002-2016, (see Equation 1), for the IW estimator by Sun and Abraham (2021). Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

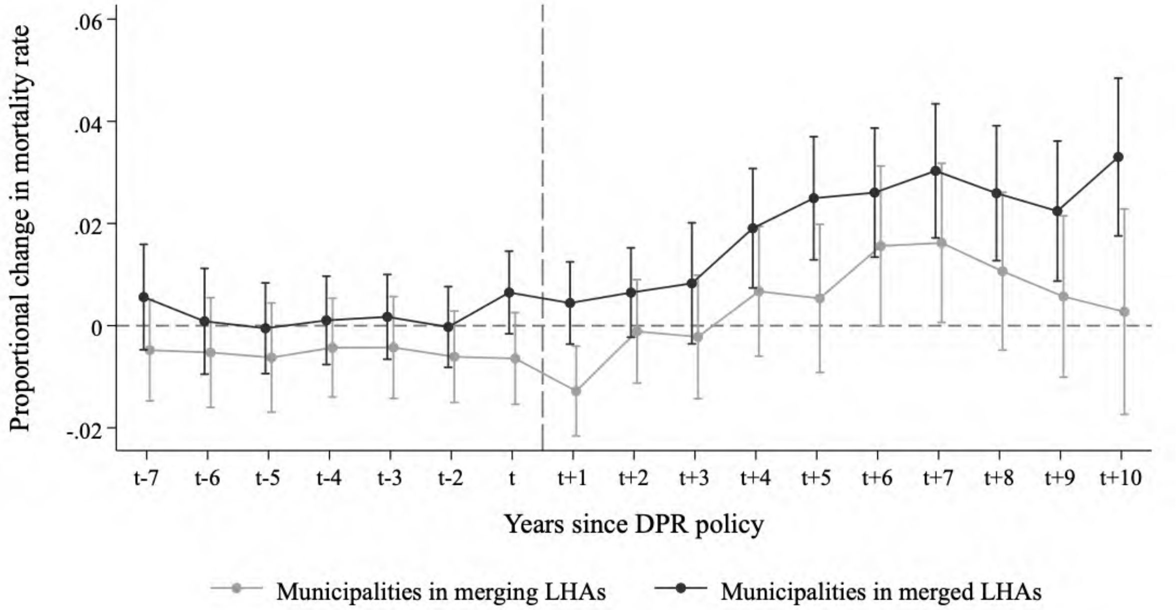
units as controls with estimates from the sample restricted to use last-treated units as controls. The plot reveals systematically larger coefficients in the latter case. The policy begins to affect mortality at $t + 1$, marked by a 1% increase in the mortality rate, and reaches a plateau five to seven years since the mergers, with an increase of approximately 3.5% relative to the baseline average of 985.88 deaths per 100,000 residents. If our hypotheses about the higher degree of similarity are correct, then these findings can be considered as an upper bound of the main effect estimated when comparing all treated municipalities to never-treated municipalities.

6.5 Heterogeneity by municipality type

We consider whether the overall mortality increase estimated in all municipalities varies with municipal characteristics. In particular, to detect potential heterogeneity patterns behind the overall effect, we classify Italian municipalities and the LHAs to which they belong according to several dimensions.

As a first exercise, we explore the heterogeneity in the effects of consolidation by distinguish-

Figure 11: Effects of LHAs consolidation on mortality rate by merger status



Notes: Dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018, (see Equation 1), for the IW estimator by Sun and Abraham (2021). The dependent variable is the inverse hyperbolic sine of the population mortality rate per 100,000 residents calculated for all ages and causes in different groups of municipalities. All never-treated municipalities are used as controls. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

ing municipalities based on whether their LHAs retained the same headquarters (*merging*) or not (*merged* or absorbed LHAs). Proximity to the headquarters has been emphasized by the literature at the intersection of decentralization and development as an important factor because it affects the quality and flow of information regarding local needs (Dahis and Szerman, 2024; Oates, 1999). Figure 11 reveals that municipalities within merging LHAs do not follow a definitive trajectory of population health outcomes, either harmful or beneficial. While a favorable effect is observed in the initial year following the merger, indicated by a 1.3% reduction in mortality relative to a baseline rate of 997 deaths per 100,000, this improvement is not sustained over time. In parallel, a weak positive effect on mortality, comparable in magnitude (+1.6%), is observed six to seven years post-merger but similarly lacks persistence. Not surprisingly, the ATT estimates are not significantly different from zero. In contrast, the impact of the consolidation policy on mortality is statistically significant and adverse for municipalities belonging to merged LHAs. Consistent with our main results, a statistically significant increase in mortality is observed, with an ATT of 1.7%, which rises to 2.5% when focusing on the restricted post-merger window $[t + 4,\ t + 9]$, relative to a baseline mortality rate of 989.5 deaths per 100,000 residents.

We now focus on differences between municipalities of varying population sizes and those belonging to LHAs with different dimensions as measured by the population size of their catchment area. As shown in Table 1, municipal size only entails weak differences. Specifically, the first two columns of Table 1 present event-study coefficients separately for small and large municipalities, with municipal size defined relative to the median population in 2002 (approximately 2,759 inhabitants). The first column provides clear evidence of long-term deterioration of population health for smaller municipalities, which already had a higher average baseline mortality rate (1,153 deaths per 100,000 residents in the year before the consolidation). Statistically significant effects only appear from year $t+7$ onward, indicating a 3.5% increase in the mortality rate. For larger municipalities, as shown in the second column, the negative impact is observed earlier, emerging four years after the reform and peaking at $t+6$ with a 2.2% increase in the mortality rate relative to a lower average of approximately 975 deaths per 100,000 residents. However, the overall ATT in the $[t+4,\, t+9]$ window is smaller for larger municipalities (1.7%) compared to smaller ones (2.1%). On the contrary, the largest discrepancies arise when focusing on the LHAs size. By categorizing LHAs as small or large according to the median LHA population (529,588 residents in 2018), we can see, in Columns (3) and (4), that the negative effect of the policy only occurred in large LHAs.

Potential additional challenges in reorganizing health care services may arise when a sharp increase in the population under the responsibility of a given LHA occurs post-merger. To explore this hypothesis, we categorize treated municipalities based on whether their LHA experienced a smaller or larger population increase post-merger, dividing them into two groups: those above or below the median change in population size. The last two columns of Table 1 present event-study estimates that, differently from the previous 4 columns, are obtained by comparing treated municipalities within each group to all never-treated municipalities (those within LHAs that did not undergo similar substantial population changes). The results show that the mortality increase is concentrated in the group of municipalities that underwent a major increase in the catchment area of their LHAs, with the ATT in the restricted effect window reaching 2.9%. Importantly, the two groups have comparable baseline mortality rates, ruling out potential convergence effects that could be expected if starting points were significantly different. Overall, these findings suggest that economies of scale, a potential benefit arising from consolidation, do not manifest in improved health outcomes, at least when considering the

Table 1: Effects of LHAs consolidation on mortality rate by municipality size, LHA size and LHA size increase.

	(1) Small municipality	(2) Large municipality	(3) Small LHA	(4) Large LHA	(5) Small LHA increase	(6) Large LHA increase
t-7	0.006	0.001	0.012	0.003	0.011^{*}	-0.004
	(0.011)	(0.004)	(0.007)	(0.006)	(0.006)	(0.005)
t-6	-0.006	-0.002	-0.004	-0.002	-0.002	-0.003
	(0.010)	(0.004)	(0.007)	(0.006)	(0.006)	(0.005)
t-5	-0.007	-0.003	0.004	-0.008	0.006	-0.009^{*}
	(0.010)	(0.004)	(0.007)	(0.005)	(0.006)	(0.005)
t-4	-0.007	-0.001	-0.004	-0.003	0.004	-0.005
	(0.009)	(0.004)	(0.006)	(0.005)	(0.006)	(0.004)
t-3	-0.002	-0.001	-0.001	-0.001	0.004	-0.004
	(0.009)	(0.004)	(0.006)	(0.005)	(0.005)	(0.004)
t-2	0.003	-0.003	-0.002	-0.003	0.003	-0.006
	(0.009)	(0.003)	(0.006)	(0.004)	(0.005)	(0.004)
t	0.017^{*}	-0.002	-0.004	0.001	0.002	-0.000
	(0.009)	(0.003)	(0.006)	(0.004)	(0.005)	(0.004)
t+1	0.002	-0.005	-0.002	-0.006	-0.006	-0.003
	(0.009)	(0.003)	(0.006)	(0.004)	(0.005)	(0.004)
t+2	-0.001	0.004	-0.012^{*}	0.006	-0.005	0.007
	(0.010)	(0.004)	(0.007)	(0.004)	(0.006)	(0.005)
t+3	-0.006	0.005	-0.004	0.004	-0.000	0.006
	(0.011)	(0.005)	(0.007)	(0.006)	(0.006)	(0.006)
t+4	0.007	0.014^{***}	-0.008	0.019^{***}	-0.001	0.022^{***}
	(0.012)	(0.005)	(0.008)	(0.006)	(0.007)	(0.006)
t+5	0.007	0.016^{***}	-0.005	0.022^{***}	-0.000	0.027^{***}
	(0.012)	(0.006)	(0.008)	(0.007)	(0.007)	(0.007)
t+6	0.014	0.022^{***}	0.005	0.024^{***}	0.007	0.030^{***}
	(0.013)	(0.006)	(0.009)	(0.007)	(0.008)	(0.007)
t+7	0.035^{***}	0.021^{***}	0.001	0.029^{***}	0.004	0.036^{***}
	(0.012)	(0.006)	(0.009)	(0.008)	(0.008)	(0.007)
t+8	0.027^{**}	0.018^{***}	-0.009	0.026^{***}	-0.003	0.032^{***}
	(0.013)	(0.006)	(0.008)	(0.008)	(0.008)	(0.007)
t+9	0.035^{***}	0.012^{**}	-0.007	0.019^{**}	-0.004	0.025^{***}
	(0.013)	(0.006)	(0.009)	(0.008)	(0.009)	(0.007)
t+10	0.032^{**}	0.019^{**}	-0.020^{*}	0.022^{**}	0.001	0.028^{***}
	(0.016)	(0.008)	(0.011)	(0.010)	(0.010)	(0.009)
ATT	0.014	0.011^{***}	-0.004	0.014^{***}	-0.001	0.018^{***}
	(0.008)	(0.004)	(0.006)	(0.005)	(0.005)	(0.004)
$\text{ATT}_{(t+4;t+9)}$	0.021^{**}	0.017^{***}	-0.004	0.023^{***}	0.001	0.029^{***}
	(0.010)	(0.005)	(0.007)	(0.006)	(0.006)	(0.006)
Baseline average	1153.019	974.956	1048.651	975.278	1024.048	1021.388
R^2	0.571	0.852	0.732	0.819	0.767	0.809
Observations	56491	56491	56763	56219	79101	78965

Notes: Regressions are estimated using data from 2002-2018. The IW estimator by Sun and Abraham (2021) is used for all models. The dependent variable is the inverse hyperbolic sine of the population mortality rate per 100,000 residents calculated for all ages and causes in different groups of municipalities. All never-treated municipalities are used as controls in columns (5) and (6). See the note of Figure 5 for details. Standard errors are clustered at the municipality level. * p<0.10, ** p< 0.05, *** p<0.01.

very large sizes sometimes reached in Italy.[35] This reinforces the evidence that municipalities in significantly large LHAs incurred considerable health costs, as reflected in higher mortality rates. Factors such as increased bureaucratic complexity, reduced patient-provider proximity,

[35]Note that, to obtain a broad indication of an optimal size, we also considered quartiles of LHAs' increases. However, the results did not change qualitatively: the smallest two quartiles are statistically unaffected by mergers, whereas the largest two quartiles display mortality increases.

Table 2: Effects of LHAs consolidation on mortality rate by municipal type.

	(1) Core	(2) Outlying	(3) Intermediate	(4) Peripheral	(5) Very perypheral
t-7	-0.010	0.003	0.012	0.018	0.058
	(0.008)	(0.006)	(0.009)	(0.012)	(0.038)
t-6	-0.008	-0.002	0.005	-0.006	0.087***
	(0.008)	(0.006)	(0.010)	(0.011)	(0.033)
t-5	-0.011	-0.001	-0.000	0.003	0.031
	(0.008)	(0.005)	(0.009)	(0.011)	(0.026)
t-4	-0.006	-0.004	0.010	0.004	0.028
	(0.007)	(0.005)	(0.009)	(0.010)	(0.028)
t-3	-0.000	-0.004	-0.003	0.007	0.062***
	(0.007)	(0.005)	(0.009)	(0.010)	(0.023)
t-2	-0.005	-0.000	-0.006	0.001	0.029
	(0.006)	(0.005)	(0.008)	(0.010)	(0.026)
t	-0.002	-0.000	0.009	0.002	0.046*
	(0.006)	(0.005)	(0.008)	(0.010)	(0.024)
t+1	-0.007	-0.001	-0.003	-0.004	0.024
	(0.006)	(0.005)	(0.008)	(0.010)	(0.025)
t+2	-0.003	0.010*	-0.002	0.001	-0.002
	(0.008)	(0.005)	(0.009)	(0.010)	(0.026)
t+3	0.004	0.010	0.002	-0.009	0.049
	(0.008)	(0.007)	(0.011)	(0.013)	(0.034)
t+4	0.005	0.022***	0.013	0.014	0.020
	(0.008)	(0.007)	(0.010)	(0.014)	(0.031)
t+5	-0.003	0.036***	0.001	0.015	0.068*
	(0.011)	(0.007)	(0.011)	(0.013)	(0.039)
t+6	0.005	0.035***	0.009	0.032**	0.045
	(0.010)	(0.008)	(0.011)	(0.014)	(0.033)
t+7	0.011	0.034***	0.028**	0.029**	0.005
	(0.012)	(0.008)	(0.012)	(0.014)	(0.037)
t+8	0.007	0.030***	0.015	0.025*	0.046
	(0.012)	(0.008)	(0.012)	(0.014)	(0.034)
t+9	0.001	0.024***	0.021*	0.019	0.031
	(0.011)	(0.008)	(0.012)	(0.015)	(0.037)
t+10	0.033***	0.013	0.031**	0.013	0.026
	(0.011)	(0.010)	(0.013)	(0.020)	(0.055)
ATT	0.002	0.020***	0.009	0.012	0.033
	(0.006)	(0.005)	(0.008)	(0.010)	(0.023)
$ATT_{(t+4;t+9)}$	0.004	0.030***	0.014	0.022*	0.036
	(0.008)	(0.006)	(0.009)	(0.011)	(0.028)
Baseline average	1070.730	910.712	1046.027	1083.127	1038.621
R^2	0.945	0.764	0.729	0.711	0.671
Observations	3638	57409	27251	19737	4947

Notes: Regressions are estimated using data from 2002-2018. The IW estimator by Sun and Abraham (2021) is used for all models. The dependent variable is the inverse hyperbolic sine of the population mortality rate per 100,000 residents calculated for all ages and causes in different groups of municipalities. See the note of Figure 5 for details. Standard errors are clustered at the municipality level. * p<0.10, ** p< 0.05, *** p<0.01

or stretched resources could explain why larger LHAs fail to achieve economies of scale, and ultimately benefit population health.

We finally extend the analysis to examine whether the policy's effects exhibit heterogeneous spatial and geographical patterns, particularly in relation to remoteness. Recent studies highlight persistent health disparities between urban and rural areas (see, e.g., Cosby et al., 2019). Evidence suggests that rural hospital closures exacerbate adverse health outcomes (Carroll, 2023), and can generate negative spillovers in neighboring urban areas (Gujral and Basu, 2019). To assess the role of remoteness in mitigating or amplifying the effects of the consolidation process, we estimate equation (1) separately for Core, Outlying, Intermediate, Peripheral, and Very Peripheral municipalities. Core municipalities serve as hubs or inter-municipal hubs, providing all three essential public services. In terms of healthcare, at least one hospital with an emergency department is required. Outlying municipalities are those with travel times to the nearest core municipality below the median of actual average travel times. Intermediate municipalities have travel times between the median and the third quartile, while peripheral municipalities fall between the third quartile and the 95th percentile. Finally, Very Peripheral municipalities are those with travel times exceeding the 95th percentile from the nearest hub. Table 2 shows that the category most adversely affected by the merger is that of Outlying treated municipalities. All the time periods since t+4 are statistically significant, determining an ATT of 2% in the post-treatment period, rising to 3% in the period $[t+4,\, t+9]$. For Intermediate municipalities, a weaker effect is observed at $t+7$, but this effect does not persist over time. Peripheral municipalities, with a higher baseline mortality rate of 1,083 deaths per 100,000 residents, exhibit significant consolidation effects beginning at $t+6$. However, these effects keep their statistical significance only for three years, resulting in an ATT of 2.2% in the restricted effect window. For Very peripheral municipalities, results are inconclusive, probably because of the reduced sample size: no overall statistical significance is found, but the coefficients' sign always points to a mortality rate increase. Overall, the latter two areas - traditionally those receiving limited public health investment and exposed to access barriers - did not benefit from being merged into larger administrative jurisdictions.

To further investigate the observed disparities, we conduct separate estimations for each municipal type (Core, Outlying, Intermediate, Peripheral, and Very Peripheral), considering the size variation in the LHA catchment area and the merger status. Although the sample size is quite reduced for a few categories, this approach allows us to better understand how the latest results, especially for Outlying municipalities, correlate with our previous analysis, which identified the most adverse outcomes for municipalities within LHAs that have experienced substantial population growth or are part of merged LHAs. Figure A4 in the Appendix generally confirms that the mortality increase mainly affects municipalities that experienced the largest population growth for their LHA following the merger. The largest effect is observed in graph (b) for Outlying municipalities, where the mortality rate increases persistently from $t+2$ onward.

The prominence of the negative health impacts for Outlying municipalities is further confirmed by Figure A5. Municipalities in this category are adversely affected by the policy regardless of their merger status, with large effects observed from $t + 5$ onward, even for municipalities belonging to merging municipalities. This contrasts with the general trend when considering all municipalities in merging LHAs, as reported in Figure 11. Overall, the consolidation process appears to have inadvertently nurtured a new frail area, potentially because of an inadequate allocation of physical and human resources within the newly expanded administrative entities.

7 Discussion and Conclusions

Over the past few decades, the Italian NHS has undergone a progressive consolidation of its health authorities, a reform aimed at addressing rising healthcare costs and enhancing efficiency by centralizing administrative functions and expanding the scale of health service provision. Similar to districts, departments, or regional units in other high-income countries, these authorities are responsible for delivering healthcare services at the local level. The consolidation process resulted in larger entities with broader catchment areas. The staggered implementation of the reform across regions has created significant variation in exposure to the policy, providing a unique opportunity to assess its impact on population health.

This study uses administrative data from the Italian Ministry of Health and municipal-level death records from 2002 to 2018. This allowed us to reconstruct the timeline of the LHA amalgamation and track municipalities' transitions to new administrative configurations. We exploit this data within an event-study DID approach to estimate the policy impact on mortality rates in municipalities exposed to LHA consolidation.

Our analysis reveals a significant increase in mortality rates following the LHAs mergers. The policy's effect is not immediate, generally emerging four years after implementation and persisting over time. When focusing on a restricted observation window beginning four years post-consolidation and extending to nine years post-consolidation, we estimate a positive ATT of 1.8% on the total mortality rate. A back-of-the-envelope calculation suggests this corresponds to 16,027 excess deaths in treated municipalities.

We document that this effect is primarily driven by two major causes of death: diseases of the circulatory system and cancers. Moreover, the age-specific analysis highlights two concerning aspects. First, the working-age population is more affected than the oldest age group, challenging the premise that older populations bear the brunt of health system inefficiencies. Second, deaths attributable to conditions generally considered preventable disproportionately contribute to the overall variation in mortality among individuals aged 0–74.

The adverse health effects were not evenly distributed. A key finding is that the negative effect on mortality was predominantly observed in larger LHAs, and those experiencing a significant expansion of their catchment area. As highlighted in Section 1, the Italian reform created extremely large local entities compared to other countries. While the primary objective was cost containment, consolidation was also expected to improve service quality by centralizing outpatient healthcare services. The observed mortality penalty in larger LHAs suggests that other factors may have outweighed those benefits. These include failure to manage increased organizational complexity and address the unique characteristics and diverse health needs of different areas and population groups.

An additional finding is that the rise in mortality is largely confined to municipalities of the merged LHAs, while those of merging LHAs generally remain unaffected. One initial interpretation is that this outcome stems from the headquarters of the LHAs prioritizing areas they managed prior to consolidation. However, a more detailed analysis, considering the spatial distribution of essential public services, reveals that the most affected municipalities are those classified as Outlying – located in the surrounding areas of Core municipalities where services are concentrated. Notably, in Outlying areas, mortality increases significantly even for municipalities in the merging LHAs. The policy resulted in 10,705 additional deaths only in Outlying areas. Understanding the mechanism behind the creation of these new vulnerabilities warrants further investigation, particularly to understand how the amalgamations have altered the allocation of physical and organizational resources within the newly expanded administrative entities.

To better understand the mechanism behind our results, future research would benefit from access to individual-level data. Such data could help determine whether, for example, larger LHAs failed to promote adequate prevention, earlier diagnosis, treatment compliance, or equitable access to healthcare services. On the service provision side, examining data on healthcare professionals' mobility across municipalities after amalgamations could provide valuable evidence on whether consolidations led to a redistribution of professionals, favoring hospitals and clinics located in core areas at the expense of surrounding areas. While we do not consider our findings conclusive, the absence of positive effects on mortality was unexpected, particularly in the long run and for specific categories of municipalities. Nevertheless, our analysis raises concerns about hastily applying a "bigger is better" policy principle to healthcare administration. Policymakers should exercise caution before adopting consolidation policies. A more comprehensive approach is essential, one that carefully evaluates the impact on healthcare quality and considers potential cost reductions alongside the risk of unintended consequences for population health.

Declaration of generative AI and AI-assisted technologies in the writing process.

Statement: During the preparation of this work the authors used Microsoft CoPilot in order to improve the readability and language of the manuscript. After using this tool, the authors reviewed and edited the content as needed and take full responsibility for the content of the published article.

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Appendix

Table A1: Summary statistics by consolidation cohort.

	Cohort 1		Cohort 2		Cohort 3		Cohort 4		Cohort 5		Cohort 6		Cohort 7		Cohort 8		Cohort 9		Cohort 10		Cohort 11		Cohort 12		Treated	Never treated
	pre	post	pre	post	pre	post	pre	post	pre	post	pre	post	pre	post	pre	post	pre	post	pre	post	pre	post	pre	post		
Mortality rates																										
All ages, all causes	1,087.8	1,117.9	1,157.4	1,218.0	833.1	907.0	1,206.1	1,239.1	990.7	1,095.5	1,246.2	1,324.9	1,141.0	1,206.8	1,211.0	1,230.5	1,083.2	1,107.4	1,182.1	1,286.0	961.6	1,084.3	965.4	1,039.2	1,084.3	1,114.5
Age 0-14	34.4	23.8	33.8	24.3	37.9	28.1	41.9	24.8	35.1	28.9	41.1	28.2	39.7	32.4	30.0	20.6	31.5	16.1	29.5	20.1	28.9	21.3	28.2	18.3	28.9	31.2
Age 15-49	92.6	76.0	79.7	72.4	78.8	66.3	92.0	77.3	90.2	75.9	90.7	74.9	82.4	65.6	78.6	60.5	77.6	62.6	81.0	62.8	76.4	60.2	75.6	60.8	76.6	77.9
Age 50-64	517.7	442.4	479.9	422.4	483.5	398.7	537.2	466.3	554.7	487.8	509.1	444.0	557.1	458.3	455.9	358.1	432.6	371.1	568.4	453.7	468.8	386.1	463.3	366.0	469.7	480.9
Age 65+	4,418.4	4,372.7	4,292.2	4,456.2	4,076.3	4,010.4	4,632.2	4,570.7	4,381.8	4,576.0	4,470.2	4,753.5	4,522.5	4,729.4	4,374.7	4,311.1	4,294.5	4,245.4	4,386.7	4,277.9	4,179.7	4,067.1	4,171.9	4,064.9	4,330.0	4,406.1
Cancer	324.2	327.7	289.5	301.1	230.5	248.6	299.5	308.7	239.1	261.6	284.0	286.5	225.5	215.3	323.8	311.4	326.2	307.0	367.0	377.8	282.2	319.6	282.4	291.1	290.8	299.1
Circulatory diseases	448.9	406.0	517.1	499.1	337.9	339.5	511.2	481.7	444.9	475.5	545.9	558.9	545.6	564.6	508.0	459.5	424.4	391.1	427.9	432.5	364.8	365.5	364.3	353.4	428.0	433.2
Respiratory diseases	77.1	91.2	77.5	85.1	60.5	70.5	82.9	88.6	66.5	75.5	83.2	95.2	71.6	90.9	83.9	101.2	79.9	99.9	84.6	110.4	63.7	84.6	64.1	81.6	76.1	75.5
Amenable causes	97.2	77.7	93.8	73.6	76.4	64.4	101.2	86.5	94.3	82.1	95.0	78.8	86.7	84.7	84.8	66.7	69.1	56.5	93.5	84.2	72.8	65.2	72.1	62.8	79.8	83.9
Non-amenable causes	271.2	238.4	254.4	221.7	218.8	197.7	293.4	258.1	257.5	234.4	263.8	240.3	260.5	223.2	254.3	212.6	233.8	206.8	315.7	297.9	236.3	225.6	235.3	215.7	244.9	254.5
Municipalities characteristics																										
Per capita income	13,583.7	15,266.3	8,158.3	10,062.4	8,043.2	10,717.6	9,554.1	11,526.9	5,921.8	7,885.2	7,820.3	10,014.9	5,623.1	7,761.6	9,854.8	11,770.3	11,483.2	13,421.9	12,103.3	14,253.9	11,150.8	14,267.0	11,279.6	13,556.5	10,682.8	10,914.5
% pop. age 0-14	12.4	13.5	13.1	12.4	16.8	15.1	12.9	12.6	15.8	13.6	12.3	11.5	15.1	13.6	12.3	12.2	13.1	13.7	11.7	11.5	14.1	13.2	14.0	13.3	13.5	13.4
% pop. age 15-49	46.6	43.7	45.6	43.4	49.3	46.6	45.5	42.6	48.3	46.0	45.0	42.4	46.3	44.0	43.9	40.6	45.9	42.3	43.9	39.0	47.3	41.6	47.0	42.0	45.2	44.9
% pop. age 50-64	19.8	20.4	17.3	19.8	16.5	18.7	19.1	20.8	16.3	19.6	17.9	21.1	16.6	19.6	19.1	21.2	18.8	20.8	21.1	22.6	19.0	21.7	19.2	22.2	19.5	19.6
% pop. age 65+	21.2	22.5	24.0	24.5	17.4	19.6	22.6	24.0	19.6	20.8	24.8	25.1	22.1	22.8	24.7	26.0	22.1	23.3	23.2	26.9	19.6	23.6	19.8	22.5	21.8	22.0
Observations	92	690	1280	4160	1510	3624	4242	7777	4109	5870	1312	1476	360	320	968	528	780	325	1183	364	8022	3033	8595	1146	67898	45084

Notes: Cohort 1 includes municipalities exposed to the LHA merger in 2004; Cohort 2 in 2006; Cohort 3 in 2007; Cohort 4 in 2008; Cohort 5 in 2009; Cohort 6 in 2010; Cohort 7 in 2011; Cohort 8 in 2013; Cohort 9 in 2014; Cohort 10 in 2015; Cohort 11 in 2016; and Cohort 12 in 2017. For each cohort, the (unweighted) Pre and post-means are calculated by averaging municipal-level indicators for the years before and after the respective LHA mergers. In the last column, the (unweighted) means for municipalities never exposed to mergers are calculated over the entire period from 2002 to 2018.

Table A2: Evolution of the LHAs consolidation process since 2002.

	2002	2003	2004	2005	2006	2007	2008	2009	2010	2011	2012	2013	2014	2015	2016	2017	2018
Piemonte	22	22	22	22	22	22	13	13	13	13	13	13	13	13	13	12	12
Valle d'Aosta	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1
Lombardia	15	15	15	15	15	15	15	15	15	15	15	15	15	8	8	8	8
PA Trento	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1
PA Bolzano	4	4	4	4	4	1	1	1	1	1	1	1	1	1	1	1	1
Veneto	21	21	21	21	21	21	21	21	21	21	21	21	21	21	9	9	9
Friuli-Venezia Giulia	6	6	6	6	6	6	6	6	6	6	6	6	5	5	5	5	5
Liguria	5	5	5	5	5	5	5	5	5	5	5	5	5	5	5	5	5
Emilia-Romagna	13	13	11	11	11	11	11	11	11	11	11	11	8	8	8	8	8
Toscana	12	12	12	12	12	12	12	12	12	12	12	12	12	3	3	3	3
Umbria	4	4	4	4	4	4	4	4	4	4	4	2	2	2	2	2	2
Marche	13	13	13	13	1	1	1	1	1	1	1	1	1	1	1	1	1
Lazio	12	12	12	12	12	12	12	12	12	12	12	12	12	10	10	10	10
Abruzzo	6	6	6	6	6	6	6	6	4	4	4	4	4	4	4	4	4
Molise	4	4	4	4	1	1	1	1	1	1	1	1	1	1	1	1	1
Campania	13	13	13	13	13	13	13	7	7	7	7	7	7	7	7	7	7
Puglia	12	12	12	12	12	6	6	6	6	6	6	6	6	6	6	6	6
Basilicata	5	5	5	5	5	5	5	2	2	2	2	2	2	2	2	2	2
Calabria	11	11	11	11	11	11	6	6	6	5	5	5	5	5	5	5	5
Sicilia	9	9	9	9	9	9	9	9	9	9	9	9	9	9	9	9	9
Sardegna	8	8	8	8	8	8	8	8	8	8	8	8	8	8	1	1	1
Total	197	197	195	195	180	171	157	148	146	145	145	143	140	139	121	101	101

Notes: Our elaborations on Ministry of Health data.

Table A3: Effects of LHAs consolidation on total mortality rate

	(1) TWFE	(2) IW	(3) DID2S
t-7	0.009***	0.001	0.002***
	(0.003)	(0.004)	(0.001)
t-6	0.000	-0.002	-0.002
	(0.004)	(0.004)	(0.002)
t-5	-0.000	-0.003	-0.002
	(0.004)	(0.004)	(0.002)
t-4	0.001	-0.001	-0.001
	(0.003)	(0.003)	(0.002)
t-3	0.001	-0.001	-0.000
	(0.003)	(0.003)	(0.002)
t-2	-0.002	-0.003	-0.002
	(0.003)	(0.003)	(0.002)
t	0.001	0.001	-0.000
	(0.003)	(0.003)	(0.002)
t+1	-0.002	-0.004	-0.005*
	(0.003)	(0.003)	(0.003)
t+2	0.002	0.004	0.000
	(0.003)	(0.004)	(0.003)
t+3	0.001	0.004	0.002
	(0.004)	(0.005)	(0.004)
t+4	0.011***	0.014***	0.013***
	(0.004)	(0.005)	(0.004)
t+5	0.012***	0.016***	0.016***
	(0.004)	(0.005)	(0.005)
t+6	0.016***	0.022***	0.020***
	(0.005)	(0.005)	(0.005)
t+7	0.018***	0.024***	0.022***
	(0.005)	(0.006)	(0.005)
t+8	0.015***	0.019***	0.016***
	(0.005)	(0.006)	(0.005)
t+9	0.010*	0.015***	0.011*
	(0.005)	(0.006)	(0.006)
t+10	0.018***	0.020***	0.017**
	(0.006)	(0.007)	(0.007)
ATT	0.008***	0.011***	0.010***
	(0.003)	(0.004)	(0.003)
$ATT_{(t+4;t+9)}$	0.014***	0.018***	0.016***
	(0.004)	(0.005)	(0.004)
Baseline average	993.189	993.189	993.189
R^2	0.791	0.792	
Observations	112982	112982	112982

Notes: Regressions are estimated using data for the period 2002-2018 (see Equation 1) using the TWFE estimator, the IW estimator by Sun and Abraham (2021) and the two-stage DiD estimator (DID2S) by Gardner (2022). The dependent variable is the total mortality rate per 100,000 residents. See the note of Figure 5 for details. Standard errors are clustered at the municipality level. * p<0.10, ** p< 0.05, *** p<0.01

Table A4: Effects of LHAs consolidation on age-specific mortality rates

	(1)	(2)	(3)	(4)	(5)	(6)
	0-14	15-24	25-49	50-64	65-74	75 and older
t-7	0.033	-0.044	0.001	0.017*	0.011*	0.008*
	(0.043)	(0.051)	(0.015)	(0.009)	(0.007)	(0.004)
t-6	-0.023	-0.017	-0.003	0.008	-0.000	0.009**
	(0.049)	(0.060)	(0.018)	(0.010)	(0.008)	(0.004)
t-5	0.069	0.026	-0.008	0.018*	0.001	0.006
	(0.046)	(0.056)	(0.017)	(0.010)	(0.007)	(0.004)
t-4	0.033	0.037	0.021	0.014	0.004	0.004
	(0.046)	(0.053)	(0.016)	(0.010)	(0.007)	(0.004)
t-3	-0.040	0.033	0.006	0.006	0.005	0.005
	(0.050)	(0.055)	(0.016)	(0.010)	(0.007)	(0.004)
t-2	0.037	0.007	-0.005	0.005	0.003	0.000
	(0.045)	(0.054)	(0.016)	(0.010)	(0.007)	(0.003)
t	0.026	-0.032	-0.006	0.008	-0.004	-0.001
	(0.048)	(0.054)	(0.017)	(0.009)	(0.007)	(0.003)
t+1	-0.042	-0.041	0.006	0.009	-0.003	-0.007**
	(0.045)	(0.058)	(0.017)	(0.010)	(0.007)	(0.003)
t+2	-0.017	-0.043	-0.008	0.004	0.003	-0.003
	(0.051)	(0.057)	(0.018)	(0.011)	(0.008)	(0.004)
t+3	0.019	-0.127*	0.017	0.013	-0.003	-0.003
	(0.054)	(0.068)	(0.020)	(0.011)	(0.009)	(0.004)
t+4	-0.041	0.059	0.013	0.025**	0.008	0.006
	(0.056)	(0.068)	(0.019)	(0.012)	(0.010)	(0.004)
t+5	-0.062	0.018	0.050**	0.019	0.000	0.007
	(0.058)	(0.066)	(0.019)	(0.013)	(0.009)	(0.005)
t+6	-0.010	-0.005	0.055***	0.035***	0.011	0.009*
	(0.058)	(0.071)	(0.021)	(0.012)	(0.009)	(0.005)
t+7	-0.002	-0.181**	0.058***	0.050***	-0.000	0.015***
	(0.060)	(0.070)	(0.021)	(0.013)	(0.010)	(0.005)
t+8	-0.018	0.002	0.066***	0.045***	-0.017*	0.017***
	(0.061)	(0.070)	(0.021)	(0.013)	(0.009)	(0.005)
t+9	0.096	-0.001	0.050**	0.037***	0.001	0.005
	(0.063)	(0.075)	(0.024)	(0.014)	(0.011)	(0.006)
t+10	-0.094	-0.048	0.032	0.039***	0.012	0.021***
	(0.061)	(0.071)	(0.026)	(0.014)	(0.010)	(0.006)
ATT	-0.005	-0.035	0.030**	0.024***	-0.000	0.005
	(0.038)	(0.047)	(0.014)	(0.008)	(0.006)	(0.003)
$\text{ATT}_{(t+4;t+9)}$	-0.006	-0.018	0.049***	0.035***	0.000	0.010**
	(0.042)	(0.050)	(0.015)	(0.010)	(0.007)	(0.004)
Baseline average	27.889	31.739	83.079	431.090	1382.872	7110.305
pseudo-R^2	0.101	0.104	0.099	0.172	0.250	0.287
Observations	84915	83623	112132	112982	112982	112982

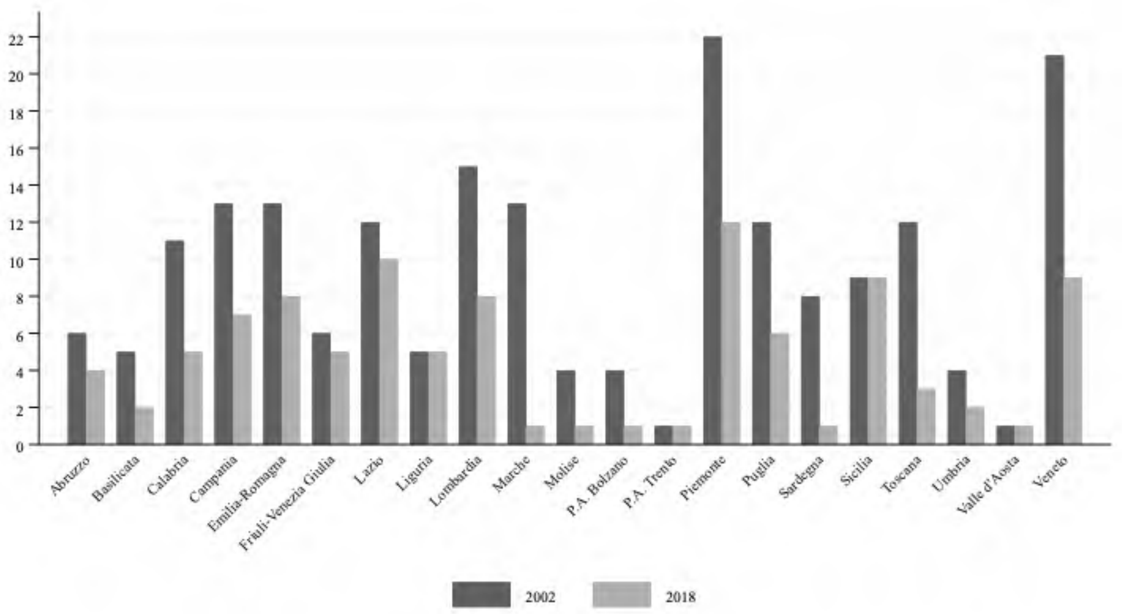
Notes: Regressions are estimated using data for the period 2002-2018 (see Equation 1). The PPML-TWFE estimator is used for all models, which include total population in 2002 as weight, municipality-fixed effects and year-fixed effects are included. The dependent variables are the age-specific mortality rates per 100,000 residents. All models include time-varying controls: per capita income and a dummy variable indicating if the municipality is in a region under the Debt Rescheduling Plan. See the note of Figure 5 for details. Standard errors are clustered at the municipality level. * $p<0.10$, ** $p< 0.05$, *** $p<0.01$

Table A5: Effects of LHAs consolidation on mortality rate by causes of death

	(1) Circulatory system diseases	(2) Cancer	(3) Digestive system diseases	(4) All other causes	(5) Amenable	(6) Non-amenable
t-7	0.022***	0.019***	0.002	-0.030***	0.018**	0.012*
	(0.006)	(0.005)	(0.012)	(0.009)	(0.009)	(0.006)
t-6	0.006	0.006	0.007	-0.015*	0.011	-0.000
	(0.006)	(0.006)	(0.012)	(0.008)	(0.011)	(0.007)
t-5	0.005	0.010*	0.016	-0.022***	0.026**	0.002
	(0.005)	(0.006)	(0.013)	(0.008)	(0.010)	(0.007)
t-4	-0.001	0.009*	0.014	-0.011	0.015	0.010
	(0.005)	(0.005)	(0.012)	(0.007)	(0.010)	(0.006)
t-3	0.004	0.002	0.018	-0.010	-0.006	0.009
	(0.005)	(0.005)	(0.011)	(0.007)	(0.010)	(0.006)
t-2	-0.004	-0.001	0.020*	-0.007	-0.002	0.006
	(0.005)	(0.005)	(0.011)	(0.006)	(0.010)	(0.006)
t	0.008*	-0.001	0.009	-0.010	0.002	-0.001
	(0.004)	(0.005)	(0.010)	(0.006)	(0.010)	(0.006)
t+1	-0.001	-0.001	-0.001	-0.005	0.010	-0.003
	(0.005)	(0.005)	(0.011)	(0.007)	(0.010)	(0.007)
t+2	0.006	0.001	-0.004	-0.006	0.005	-0.001
	(0.005)	(0.006)	(0.011)	(0.007)	(0.011)	(0.007)
t+3	0.002	0.006	0.003	-0.006	0.014	-0.000
	(0.006)	(0.007)	(0.014)	(0.008)	(0.013)	(0.008)
t+4	0.008	0.013**	0.026*	0.007	0.027*	0.011
	(0.006)	(0.006)	(0.014)	(0.009)	(0.014)	(0.008)
t+5	0.015**	0.014**	-0.001	0.004	0.018	0.012
	(0.007)	(0.007)	(0.015)	(0.009)	(0.013)	(0.008)
t+6	0.029***	0.025***	-0.030**	-0.004	0.037***	0.021**
	(0.007)	(0.007)	(0.015)	(0.010)	(0.013)	(0.008)
t+7	0.030***	0.018**	0.011	-0.001	0.046***	0.014
	(0.007)	(0.007)	(0.016)	(0.010)	(0.013)	(0.009)
t+8	0.030***	0.018**	0.003	-0.005	0.028**	0.009
	(0.008)	(0.007)	(0.016)	(0.011)	(0.014)	(0.009)
t+9	0.025***	0.010	-0.018	-0.005	0.015	0.023**
	(0.008)	(0.008)	(0.017)	(0.011)	(0.016)	(0.009)
t+10	0.024***	0.009	0.003	0.022*	0.008	0.024***
	(0.009)	(0.008)	(0.019)	(0.012)	(0.014)	(0.009)
ATT	0.015***	0.010**	-0.000	-0.003	0.020**	0.008
	(0.005)	(0.005)	(0.010)	(0.007)	(0.009)	(0.005)
$ATT_{(t+4;t+9)}$	0.023***	0.016***	-0.001	-0.000	0.028***	0.015**
	(0.006)	(0.006)	(0.012)	(0.009)	(0.010)	(0.006)
Baseline average	376.665	283.448	71.043	262.034	72.994	224.226
pseudo-R^2	0.568	0.389	0.286	0.495	0.157	0.256
Observations	112982	112982	112846	112982	112965	112982

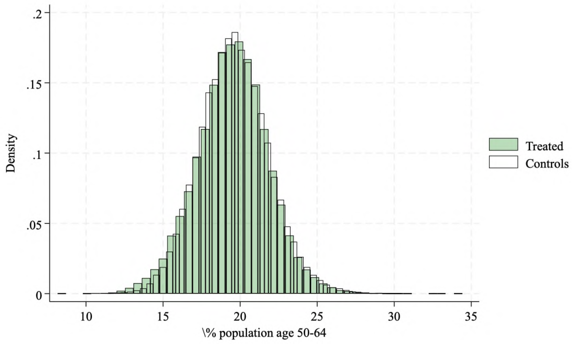
Notes: Regressions are estimated using data for the period 2002-2018 (see Equation 1). The PPML-TWFE estimator is used for all models, which include total population in 2002 as weight, municipality-fixed effects and year-fixed effects are included. The dependent variables are mortality rates per 100,000 residents calculated for the causes of death that explain more than 10% of total deaths, and for amenable and non-amenable causes of death in the population aged 0-74. The amenable and non-amenable mortality models do not include the shares of individuals aged 75-84 and 85 and older. See the note of Figure 5 for details. Standard errors are clustered at the municipality level. * p<0.10, ** p< 0.05, *** p<0.01

Figure A1: Regional breakdown of Italian Local Health Authorities configuration

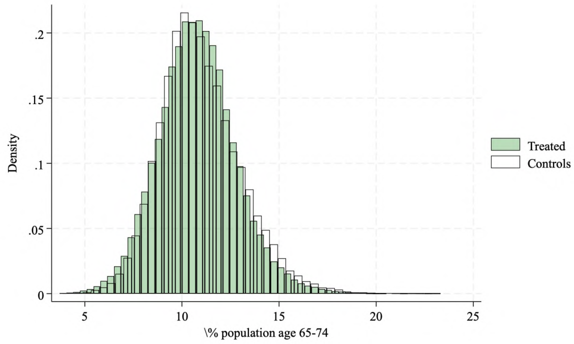


Notes: Authors' elaborations on Ministry of Health data and documental information on Local Health Authorities' consolidation.

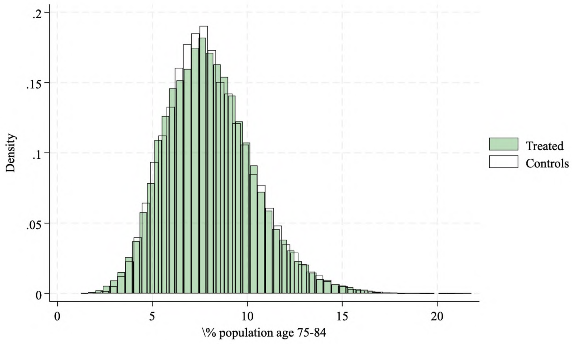
Figure A2: Distributions of covariates by treatment status in the years 2002-2018



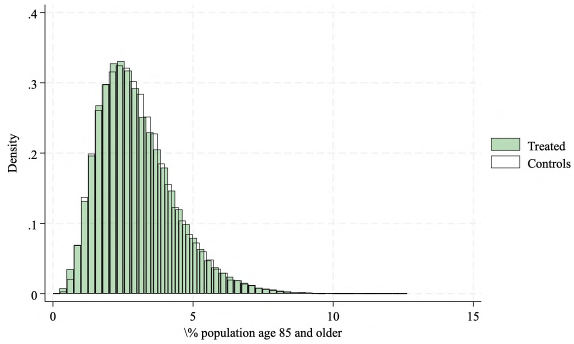
a. Population aged 50-64 (%)



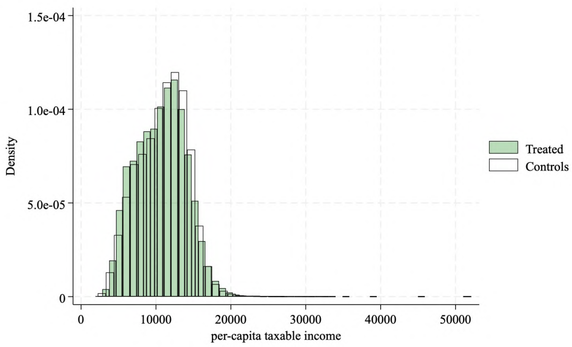
b. Population aged 65-74 (%)



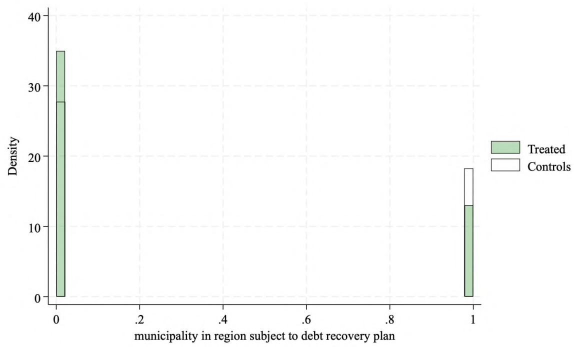
c. Population aged 75-84 (%)



d. Population aged 85 and older (%)

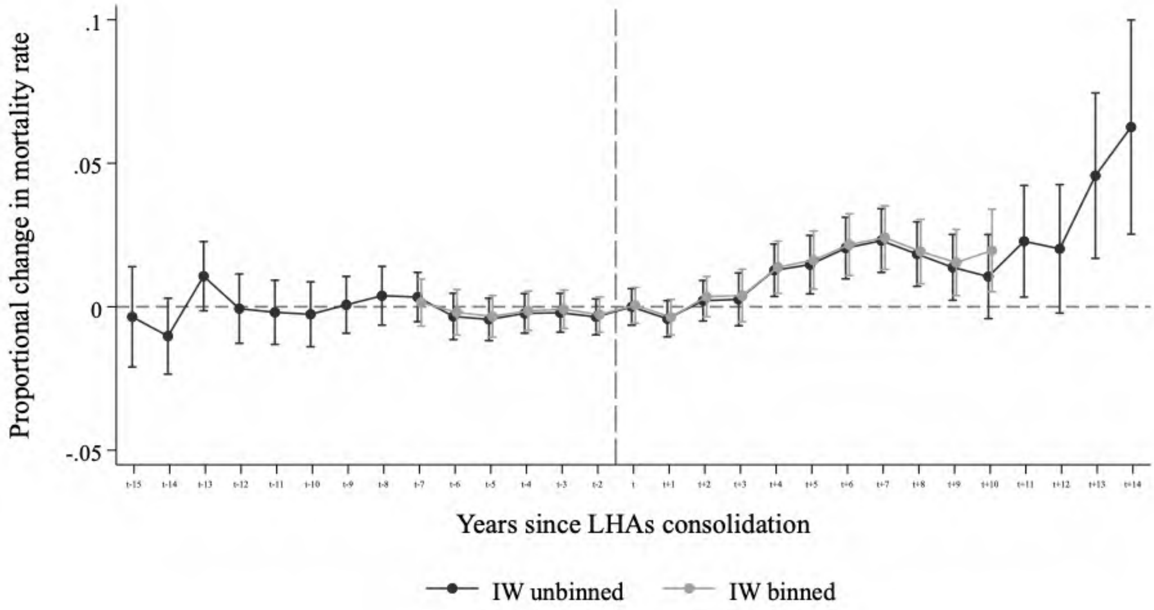


d. Per-capita taxable income



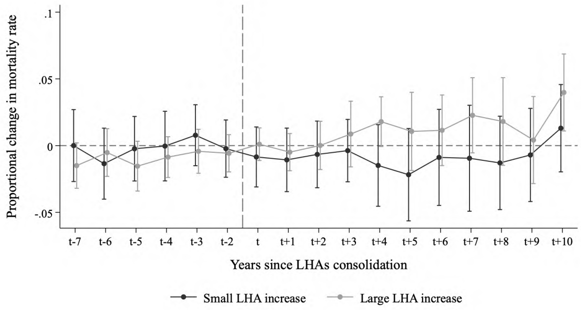
d. Under DRP

Figure A3: Comparison of binned and unbinned event-study estimates using IW estimator

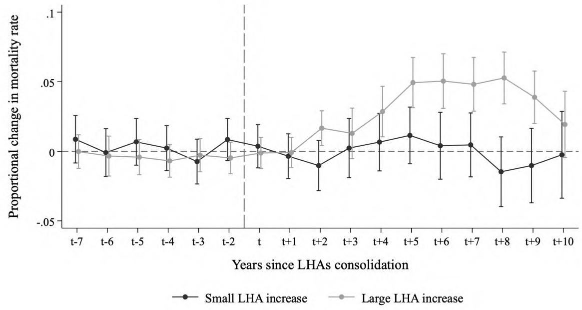


Notes: Dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018 (see Equation 1) using the IW estimator by Sun and Abraham (2021). The dependent variable is the inverse hyperbolic sine of the mortality rate per 100,000 residents calculated for all ages and causes of death. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

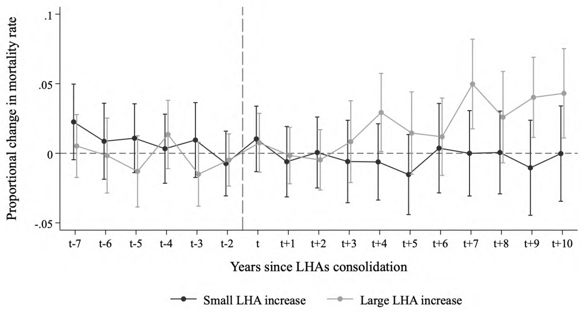
Figure A4: Heterogeneity by municipal type and LHA size increase



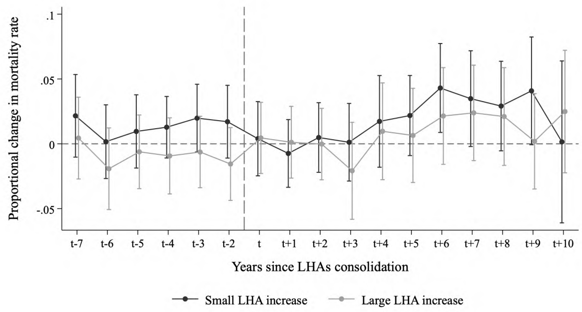
a. Core



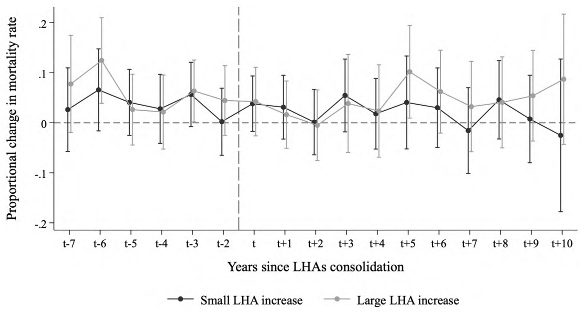
b. Outlying



c. Intermediate



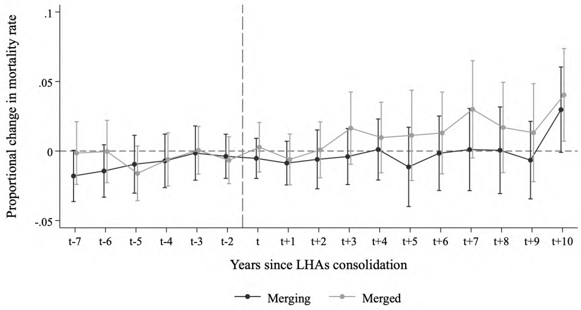
d. Periphery



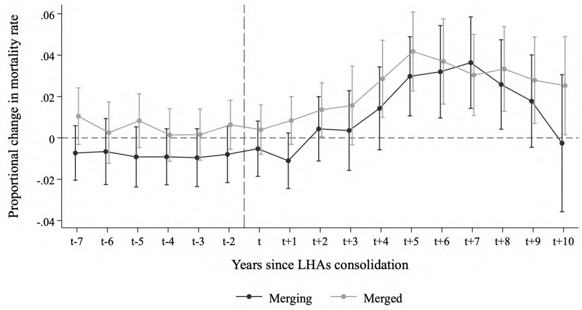
e. Very periphery

Dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018 (see Equation 1). The IW estimator by Sun and Abraham (2021) is used for all models. The dependent variable is the inverse hyperbolic sine of the population mortality rate per 100,000 residents calculated for all ages and causes in different groups of municipalities. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.

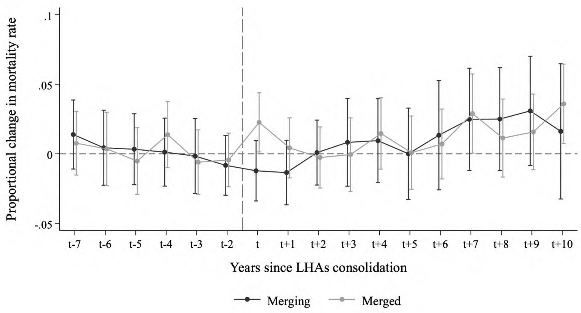
Figure A5: Heterogeneity by municipal type and merging status



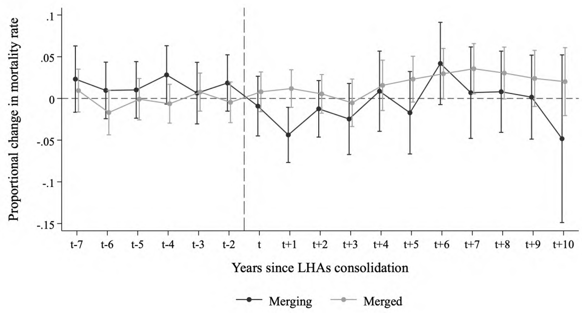
a. Core



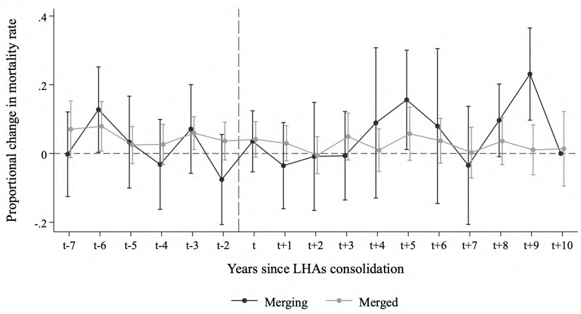
b. Outlying



c. Intermediate



d. Peripheral



e. Very peripheral

Notes: Dots indicate the estimated coefficients for each year before and after the consolidation reform using data for the period 2002-2018 (see Equation 1). The IW estimator by Sun and Abraham (2021) is used for all models. The dependent variable is the inverse hyperbolic sine of the population mortality rate per 100,000 residents calculated for all ages and causes in different groups of municipalities. Confidence intervals are indicated at the 95 per cent level. See the note of Figure 5 for details.