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Mortality evolution in Italy: the end of regional convergence?

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Abstract

Between 1992 and 2007, the Italian healthcare system underwent a profound and complex transformation. This reform rekindled interest in the study of health and mortality inequalities in Italy, and fears were expressed that the new system could trigger a rise in health disparities across regions. We contribute to the debate examining the evolution of life expectancy across Italian regions from 1974 to 2019, focusing on regional convergence. Applying sigma- and beta-convergence analysis in a novel way, we detect structural breaks—marked trend variations—occurring shortly after the most significant policy reforms of the national health system, with stronger discontinuities among males. While not establishing any direct causal link, our findings suggest that the legislative changes of the period, focused on decentralizing responsibilities and management, may have halted the pre-existing trend towards homogenization and possibly opened up a new phase of re-emerging regional survival disparities.

Introduction

During the 1980s and 1990s, many European countries (Belgium, France, Portugal, Spain and the UK, to name a few) implemented a gradual transfer of functions from central to regional governments, including some that relate to the management of the health systems (Montero-Granados et al., 2007). This happened for several reasons. For instance, many countries had accumulated a large healthcare deficit, partly due to their aging populations. Transferring managerial and financial responsibility to the regions was seen as a way to improve the effectiveness of the health system, by placing it in the hands of local policymakers more aware of the characteristics and needs of local populations (Cavalieri & Ferrante, 2020). However, the regionalization of healthcare systems has also raised concerns: this type of transformation could lead to an increase in regional disparities in health and mortality. Actually, the signs of increasing regional health disparities in Europe are multiplying. For instance, a very recent analysis of the regional life expectancy convergence in 16 different European countries identifies a generalised slowdown in the second part of the period of analysis, 1995–2019 (Sauerberg et al., 2024). This aligns with the conclusions of Maynou et al. (2015) at the European level and Wilson et al. (2020) in their comparative study of Finland and Sweden.

During the 1990s and 2000s, the Italian National Healthcare Service (NHS) too underwent a profound process of restructuring and modernization (De Belvis et al., 2023; Ferrè et al., 2014; Toth, 2014; Toth et al., 2014). Launched in 1978, the NHS was originally designed to work under the direct control of the central government, but the reforms of 1992 and 2001 introduced decentralization and transferred control to regional authorities. Since then, the central government has played a relatively marginal role, allocating resources to each of the 20 regions (based on objective parameters: number of residents and, separately, old residents in each region) and determining the essential levels of care (LEA) that must be guaranteed throughout the national territory. Budget balance in the health sector proved an issue since the beginning, and some regions, especially in the south of Italy, started to incur large deficits (Aimone Gigio et al., 2018). The so-called "piani di rientro" (recovery plans), introduced in 2007, forced the most indebted regions to restore financial balance, by increasing drug prices and fees, reducing staff, and reorganizing the service, e.g., closing down small hospitals and other facilities. After the onset of the financial crisis of 2008, however, cost containment policies were extended to all regions (Egidi & Demuru, 2018; Salinari et al., 2023).

Reforms and budget cuts revived the interest towards health and mortality inequalities in Italy, which, many scholars feared, would re-start to widen after the convergence brought about by the 1978 unification of the NHS. However, research on the evolution of health disparities over the last 20 years has yielded mixed results. Nigri et al. (2022), for instance, argue that disparities have decreased. The longevity gap between men and women and between the more and the less educated has shrunk, while mortality differences between regions and macro-areas (north, center, and south) have varied little, if at all (Caselli et al., 2021). An analysis of sigma- and beta-convergence across Italian regions between 1996 and 2016 seems to confirm this generally positive picture, both for infant mortality and life expectancy in general (Cavalieri & Ferrante, 2020).

Not everybody agrees, however. Depalo (2019), for example, using a bound inference method, argues that recovery plans caused a mortality increase in four regions: Abruzzo, Campania, Lazio, and Sardinia (please note that three of them belong to the south). Arcà et al. (2020), using an instrumental variables approach, find that recovery plans increased avoidable mortality by about 3%, and the same technique leads Cirulli and Marini (2023) to identify a negative effect of recovery plans on various indicators: mortality rates, amenable and preventable deaths. Finally, Guccio et al. (2023), with a difference-in-differences analysis, find that recovery plans slowed down survival progress, especially in peripheral municipalities and most notably in those without healthcare services, and far from them.

To sum up, research on the effects of recovery plans tends to identify distinct mortality trajectories in the regions subjected to this procedure. However, not much emerges in the analyses that compare the *evolution* of regional disparities in health and mortality. These conflicting results may depend in part on the different methodologies employed for the analyses and in part on periodization differences. For instance, Cavalieri and Ferrante (2020), who focus on the post-reform period, may correctly conclude that there was no divergence in the post-reform period (which is what we find too, see below), but cannot assess whether there was a *change* in the intensity of the convergence process. This is precisely the research question we address in this paper: in the transition from the pre- to the post-reform period, is it possible to observe a slowdown in the convergence process across the various regions? To answer this question, we analyze the evolution of regional life expectancies over a longer period, from 1974 to 2019, including some pre-reform years. On these series we employ sigma and beta-convergence analysis. With the former (sigma convergence), we look at the time evolution of a dispersion index (the variance in our case). With the latter (beta-convergence), we investigate if regions starting with lower life expectancy at birth are systematically associated with higher survival improvements in subsequent years, and thus are converging with other regions. However, in our analysis it is also necessary to identify a date (a year) that marks the end of the "old regime" and the beginning of the new one. Unfortunately, in our case, the reform we are talking about (regionalization and subsequent spending cuts) spanned 15 years, from 1992 to 2007, and it is not clear, a priori, whether the dividing line is best represented by the reform of 1992, that of 2001, or the recovery plans that began in 2007. Additionally, these interventions are unlikely to generate immediate health effects: it seems reasonable to expect that regional survival will be affected with some delay, if at all. To circumvent these dating problems, we started from a methodology aimed at identifying structural change points, and applied it, we believe, in an original way, to the study of convergence of regional life expectancies. In essence, the (beta-) convergence process of Italian regions is first analyzed using short (5-year) moving time windows, thus associating every calendar year with a measure of convergence intensity. On this new series, using econometric techniques, we search for structural change points, that is years in which a significant and prolonged change in the intensity of the convergence process occurs. Our hypothesis is that such time break exists and occur (shortly) after the healthcare system reforms that we discussed earlier. To be sure, our analysis is exploratory and not causal: the breaks that we find in the post-reform years may not be caused by the healthcare reform itself. However, our findings lend support to the hypothesis that such a causal link exists and justify further (and finer) research.

The rest of this paper is structured as follows: in the next section, we detail the data and the methods used for our analysis. In the third section, we present our results, which we then discuss in the fourth section.

Data and methods

Our data come from the Italian regional life tables of 1974–2019, produced by ISTAT, the Italian National Institute of Statistics. Of these tables, we considered e_0, separately by gender. The territorial units of our analysis are the Italian administrative regions, or NUTS 2. There are 20 of them, grouped in three to five macro-areas,[1] but as Valle d'Aosta and Molise are small, we eventually decided to merge them with the neighbouring regions with which they were originally united, Piemonte and Abruzzo, respectively. In the merging process we used weighted averages, to keep population sizes into account. This left us with 18 regions, the joint evolution of whose life expectancies we studied separately with both beta- and sigma-convergence, borrowing a methodology

[1] North comprises Valle d'Aosta, Piemonte, Liguria, Lombardy (north west) plus Trentino-Alto Adige, Veneto, Friuli-Venezia Giulia, Emilia-Romagna (north east). Centre includes Toscana, Marche, Umbria, Lazio; *Mezzogiorno* encompasses the south (Abruzzo, Campania, Molise, Puglia, Basilicata, Calabria) and the two main islands (Sicilia and Sardinia). Sometimes, however, the term south is used loosely, to indicate *Mezzogiorno*.

that was first developed in economics (Barro & Sala-i-Martin, 1990, 1992; Barro et al., 1991) and later exported to other disciplines, including demography (Hrzic et al., 2021).

Sigma-convergence

Let us consider sigma-convergence first, the most intuitive approach to the study of the problem. A dispersion index is considered (for instance the standard deviation or, as in our case, the variance) and its evolution over time is observed: sigma-convergence occurs when dispersion decreases.

Let $e_{t,r}$ denote life expectancies at birth at a generic time t for a generic region r, of which we have $R=18$. Following Hrzic et al. (2021), we used weighted variance $\mathrm{WVAR}(e_t)$ as an index of dispersion:

$$\mathrm{WVAR}(e_t) = \frac{1}{\sum_{r=1}^{R} P_{t,r}} \sum_{r=1}^{R} P_{t,r} \left(e_{t,r} - \overline{e_t}\right)^2,$$

where $P_{t,r}$ and $e_{t,r}$ are, respectively, the population and the life expectancy of region r at time t, while $\overline{e_t}$ is the weighted mean of $e_{r,t}$.

While sigma-convergence gives an idea of the evolution of survival variability across regions, it misses some potentially important elements. For instance, in the extreme case of ranking reversal, with top regions going to the bottom and vice versa, dispersion could remain (almost) the same, falsely signalling that (almost) nothing changed.

Beta-convergence

To overcome this potential limitation, following Barro et al. (1991), we also analyzed the so-called beta-convergence, checking whether regions with low initial values experienced comparatively faster progress, thus catching up with the rest. The two analyses are complementary and should be carried out jointly (Janssen et al., 2016), keeping in mind that "*beta-convergence is a necessary condition for the existence of sigma-convergence, while sigma-convergence might not accompany beta-convergence*" (Gächter & Theurl, 2011:3).

Let us define the following two quantities for a generic time series $y_0, \ldots, y_T$:

$$\Delta y_t = y_t - y_{t-1}$$

and

$$\Delta_k y_t = y_t - y_{t-k}$$

with k denoting a generic time lag.

Ignoring weights, the existence of beta-convergence can be assessed by looking at the β parameter in the following model:

$$\frac{1}{T} \Delta_T e_{T,r} = \alpha + \beta e_{0,r} + \varepsilon_r, \qquad r = 1, \ldots, R \tag{1}$$

where $\frac{1}{T} \Delta_T e_{T,r}$ is the average annual increase in life expectancy between 0 (first year of observation, 1974 in our case) and $T=45$ (2019) across all regions r, and ε_r is the error term. A statistically significant value of β indicates the presence of a process of convergence (if β is negative) or divergence (if it is positive) between 0 and T. In our case, we can work with "natural" values (life expectancies), and not their logarithms, as economists frequently do, because, unlike economic variables (such as the GDP), life expectancy usually follows a linear and not an exponential time trend (Lee, 2019; Oeppen & Vaupel, 2002; White, 2002).

Given the heterogeneity of the Italian regions in terms of population size, we applied the population-weighted version of Model 1 minimising the function $\sum_{r=1}^{R} P_{0,r}\left(\frac{1}{T}\Delta_T e_{T,r} - \alpha - \beta e_{0,r}\right)^2$. We also created a second time series of beta coefficients, applying the weighted version of Model 1 to shorter (5-year) time intervals, to assess what happened within each sub period:

$$\frac{1}{k}\Delta_k e_{i,r} = \alpha + \beta_i e_{i-k,r} + \varepsilon_r, \qquad r = 1, \ldots, R; \quad i = k, \ldots, T \tag{2}$$

where $k+1$ identifies the length of the (fixed) time interval of analysis. In our case, for instance, where $k=4$ and we went back 4 years in time (e.g., from 2000 back to 1996), we worked with $(4+1)$ 5 calendar-year intervals. The i subscript to β denotes the final year of the 5-year period on which it was calculated. This leaves us with a time series consisting of $T-k+1$ partially overlapping components: $\boldsymbol{\beta} = (\beta_k, \ldots, \beta_T)$.

Beta-convergence and shift models

Identifying a structural change in a time series means determining the moment of change in one (or more) of the parameters that describe the phenomenon, within a pre-defined and explicit functional form. In the following, we will assume that our $\boldsymbol{\beta}$ series evolve over time as in a shift model, oscillating around an average c' before the break, and around a new average c'' after it, as shown in Fig. 1. This approach, one of the simplest possible, provides more robust and reliable results than attempting to find structural change points in the evolution of sigma-convergences, which

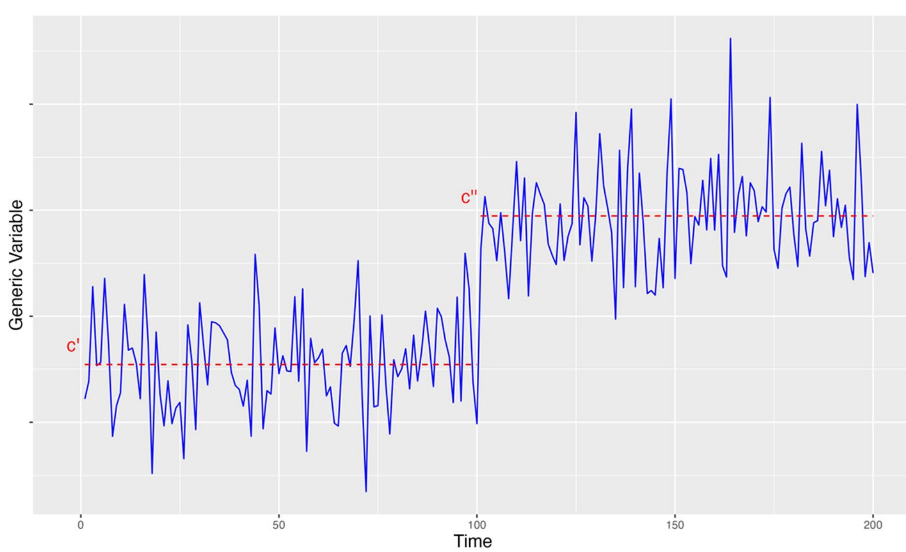

Fig. 1 Hypothetical trend of a shift model on simulated data (Source: authors' simulation)

happens to be strongly nonlinear. Therefore, when analyses based, for example, on segmented regression (as in Hrzic et al., 2021) are applied to this series, break points are detected, but their meaning is unclear. They could represent real structural change points (a change in the value of a parameter), or simply the "attempt" of the segmented regression to approximate a nonlinear trend through the juxtaposition of a succession of linear trends.

In our case, a shift unambiguously signals a change in the intensity of the convergence process. Should the sign of the average also change, this would indicate a passage from convergence to divergence (from negative to positive) or vice versa. In the simplest case, when there is only one break, our shift model can be described as follows:

$$\begin{aligned} \beta_j &= c' + \varepsilon_j; \quad j = k, \ldots, \hat{j} \\ \beta_j &= c'' + \eta_j; \quad j = \hat{j} + 1, \ldots, T. \end{aligned} \tag{3}$$

where the former equation represents the dynamics of the process up to time $\widehat{j}$ and the latter the dynamics after it, where $\widehat{j}$ is a generic possible breakpoint. Note that ε_j and η_j need not be white-noise disturbances, as it is usually assumed in standard OLS models: for instance, they could follow an ARMA process.

Following Bai and Perron (2003), we identified structural breaks by systematically analysing all possibilities, as follows. We started from a first potential breakpoint $\widehat{j}$ and we estimated the model:

$$\beta_j = \gamma_0 + \gamma_1 B_j + \omega_j, \quad j = k, \ldots, T \tag{4}$$

where:

$$B_j = \begin{cases} 1, & \text{if } j \leq \hat{j} \\ 0, & \text{if } j > \hat{j} \end{cases},$$

and where γ_0 and γ_1 are coefficients, and ω_j the error terms. The sum of squared residuals measures the goodness fit of the model pivoting around $\widehat{j}$. Next, we tried a second potential breakpoint, say $\widehat{j}'$, then a third, $\widehat{j}''$, and so on, and we proceeded until exhaustion of all possible breakpoints, repeating all the passages at each step. The breakpoints identifying a possible structural change in the series are those with the lowest sum of squared residuals.

Having identified a breakpoint, we proceeded to check whether the estimated β's behave like a shift model (Model 3 and Fig. 1) by running a unit-root test on the residuals of Model 4. If the test rejects the null hypothesis of non-stationarity, our assumption holds: before and after the breakpoint (i.e., at different levels), the series are stationary. To assess non-stationarity, we used the PP test (Perron, 1988), which incorporates heteroskedasticity.

With a recursive method, it is also possible to identify the presence of multiple structural breaks, say m, for any possible combination of sub-periods between k and T. While this is extremely time consuming, the underlying intuition remains the same. This dynamic problem can be solved in steps, by identifying first the ideal one-break partition, then the ideal two-break partition, etc. (Bai & Perron, 2003). We performed this procedure using the R library *strucchange* (Zeileis 2006; Zeileis et al., 2002, 2003). We also ran a sensitivity analysis to make sure that the choice of different values for k (i.e., using subperiods of different length) would not bias our results (see Table 2).

Results

In Italy, life expectancy is higher in the northern and central regions (except for Lazio) than in the *Mezzogiorno* (i.e., *south and islands*). In 2019, before the pandemic, life expectancy at birth in the central-northern regions ranged between 85 and 87 years and between 81 and 83 years for women and men, respectively. In contrast, in the southern regions and the islands, these values were between 1 and 2 years lower, possibly because of the higher incidence of cardiovascular diseases (Petrelli et al., 2019) and higher tobacco consumption per head (Gallus et al., 2011). In Italy, the distribution of mortality has not always followed the current north–south gradient (Caselli et al., 2021). In 1974, for instance, certain southern regions had comparatively high survival, despite their economic backwardness, e.g., in terms of per-capita income. Figure 2 shows the evolution of the spatial distribution of e_0 across Italian regions, separately by gender, in 3 years: 1974 and 2019, the first and last year of our series, and 1990, immediately before the first major reform of the Italian health system. In 1974, southern regions had a male life expectancy comparable to that of central regions and higher than that of the northern ones. Little changed until 1990, after which the north "changed pace", leaving the south behind. Regarding females, southern and northern regions had similar, and comparatively low, survival values both in 1974 and 1990, but then, just as in the case of men, northern regions improved faster. In short, the emergence of a north–south gradient in life expectancy appears to be a relatively recent phenomenon.

Figure 3 shows the evolution of life expectancy in the three traditional macro areas of Italy (north, centre, south) between 1974 and 2019, along with the weighted variance

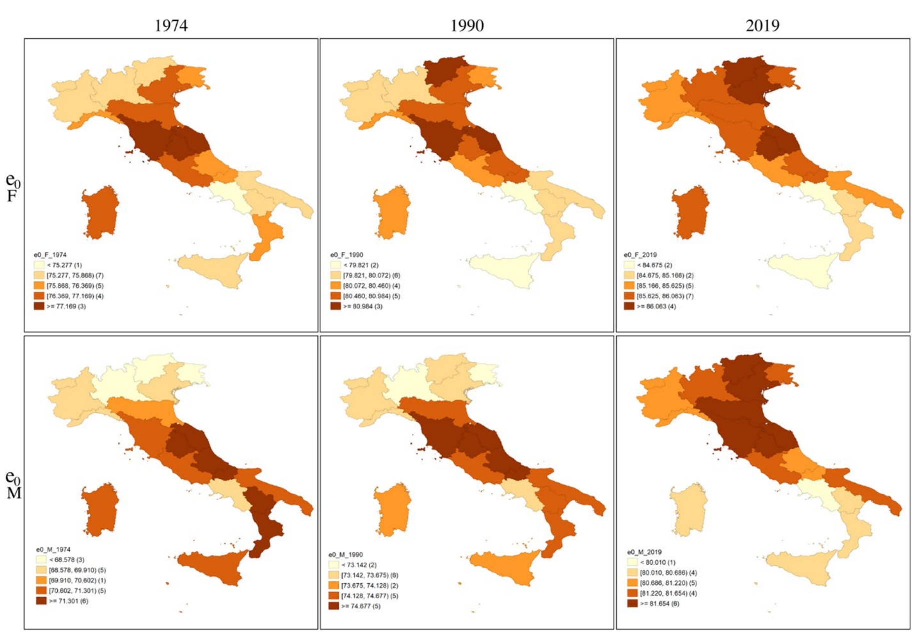

Fig. 2 Spatial distribution of e_0 by gender (F = females; M = males). Italy, selected years (Source: authors' calculations on ISTAT data (dati.istat.it))

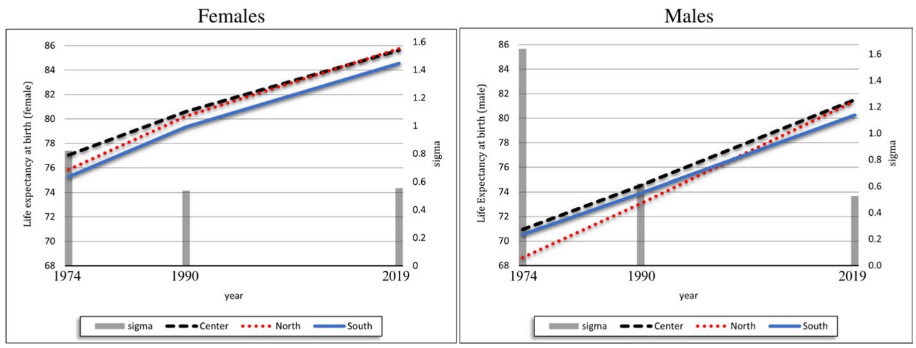

Fig. 3 Weighted averages of e_0 by macro-area (left-hand vertical scale), and sigma-convergence of e_0 (right-hand vertical scale). Italy, 1974–2019 (Source: authors' calculations on ISTAT data (dati.istat.it))

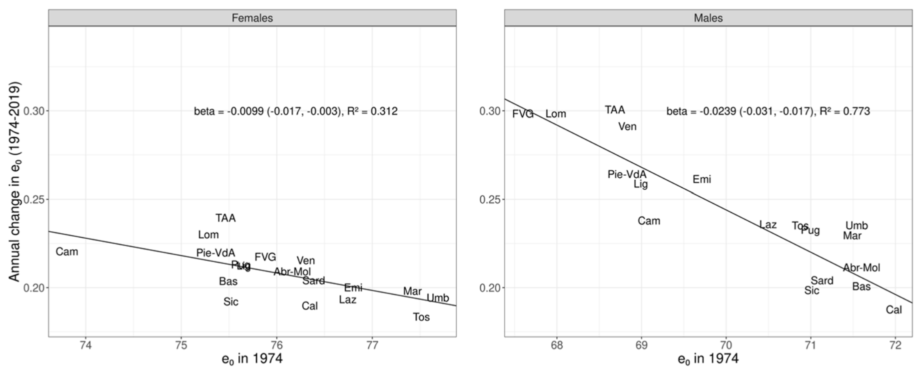

Fig. 4 Beta-convergence for regional e_0, by gender. Italy, 1974–2019. List of the region abbreviations: Abr-Mol = Abruzzo-Molise; Bas = Basilicata; Cal = Calabria; Cam = Campania; Emi = Emilia-Romagna; FVG = Friuli-Venezia Giulia; Laz = Lazio; Lig = Liguria; Lom = Lombardy; Mar = Marche; Pie-VdA = Piemonte-Valle d'Aosta; Pug = Puglia; Sard = Sardinia; Sic = Sicilia; Tos = Toscana; TAA = Trentino Alto Adige/Südtirol; Umb = Umbria; Ven = Veneto (Source: authors' calculations on ISTAT data (dati.istat.it))

of regional values, as a measure of sigma-convergence, and a special focus on 1990, our pivotal year. Between-region variability (right hand scale) strongly declined between 1974 and 1990, but barely changed afterwards. In the same period, life expectancy (left-hand scale) generally lagged behind in the south (although, initially, not for males).

To sum up, net of some differences by gender, the general message conveyed by Figs. 2 and 3 is that, in Italy, sigma-convergence slowed down after 1990, with a change in the ranking of regions, to the advantage of the northern ones.

This conclusion is reinforced by the analysis of beta-convergence, which occurs when the estimated β parameter is negative. Over the entire 1974–2019 period (Fig. 4), it did occur for both genders and especially for males, as noted also by Casacchia and Natale (2011). Things, however, are more complicated than Fig. 4 suggests: to see this, we need to break the observed interval down by sub-periods.

Figure 5 displays the evolution of the beta-convergence series estimated over 5-year intervals (Model 2), with their corresponding breakpoints. We can thus estimate an earlier break for females (the point estimate is 1993, but it is safer to consider the years of the entire 95% confidence interval, between 1990 and 2001), and a later break for males, between 1999 and 2003 (point estimate in 2001). Although we cannot locate them exactly, given the noisiness of the series, and can only indicate the period when they

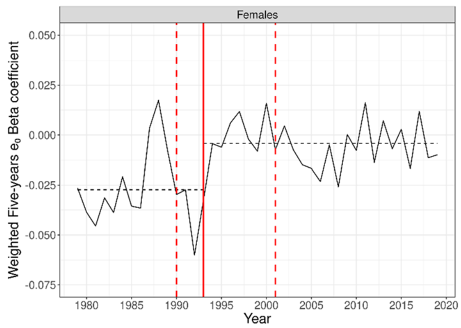

Fig. 5 Main result—time series of the 5-year beta-convergence for e_0, by gender. Vertical lines represent our estimates of the structural breaks (continuous lines: point estimates; broken lines: 95% confidence intervals). Italy, 1974–2019 (Source: authors' calculations on ISTAT data (dati.istat.it))

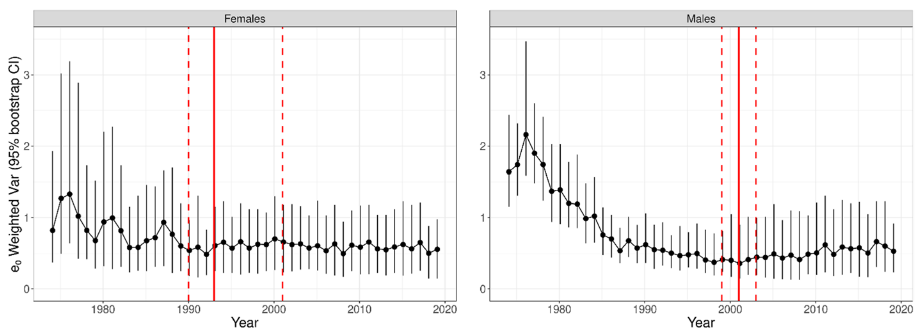

Fig. 6 Time series for sigma-convergence of e_0, differentiated by gender. The vertical lines represent the structural breaks identified in the beta-convergence time series (Source: authors' calculations on ISTAT data (dati.istat.it))

occurred, finding these breaks constitutes the main result of the present study, and, to the best of our knowledge, a novelty in this field of analysis.

Our data indicate a significant shift for males, moving from a pre-break beta-coefficient average of -0.041 to a post-break average of about 0, signalling a stop in the regional convergence process. For females, the tendency is the same, albeit a bit weaker: the shift is from a pre-break beta-convergence value of -0.027 to a post-break value of -0.004 (very close to zero, which would signal absence of convergence). Whatever happened in that period, it seems to have affected both genders, but especially men.

Figure 6 displays the sigma series together with the estimated breakpoints of Fig. 5 (i.e., those derived from the analysis of beta-convergence). Consistently with expectations, these breakpoints, although estimated on a different type of convergence and not directly applicable here, seem to separate two different phases, with convergence in the former but not in the latter.

Finally, we conducted the PP-test on the residuals of Model 4 for the beta series and in both cases the tests rejected the null hypothesis of non-stationarity. In other words, the assumptions that we made in the quest for breakpoints (Model 3) seem consistent with our data, and the beta series do behave like a shift mode (Fig. 1). Table 1 shows the results of the PP-tests for each of the variable we explored.

Table 1 Results of PP-tests launched on the residuals of Model 4, Italy, 1974–2019. Source: authors' calculations on ISTAT data (*dati.istat.it*)

Females			Males		
test-statistics	**truncation lag parameter**	***p* value**	**test-statistics**	**truncation lag parameter**	***p* value**
−32.118	3	<0.01	−41.109	3	<0.01

The null hypothesis in a PP-test is that the time series has a unit root (i.e., is non-stationary). A significantly negative test statistic suggests that the null hypothesis should be rejected (i.e., that the series *are* stationary), which one of the conditions for the analyses of this paper to be correct

The truncation lag parameter is the number of past observations (lags) considered in the stationarity-test for each current period. It is automatically chosen to correct for serial correlation, which means that the statistical test can more accurately account for patterns in the data that persist over time, thereby improving the reliability of the test's conclusion about the presence or absence of a unit root in the time series

Table 2 Structural breaks estimated in beta-convergence for different values of *k*, with the corresponding confidence intervals (sensitivity analysis), Italy, 1974–2019. Source: authors' calculations on ISTAT data (*dati.istat.it*)

k	Females		Males	
(Lag parameter)	**Phase-break**	**Point estimate**	**Phase-break**	**Point estimate**
2		–	1996–2004	1999
3	1979–1992	1985	1999–2004	2001
4	**1990–2001**	**1993**	**1999–2003**	**2001**
5	1992–2001	1995	2001–2003	2002
6	1993–2002	1995	2002–2004	2003

"–" indicates that no significant break was found

The lag parameter (variable *k*) refers to the number of years we go back to estimate the individual beta-convergence coefficients that form the time series (one time series for each value of *k*, see Model 2).The case with $k=4$ (in which we are going back 4 years, and so we are overall estimating a beta coefficient for every $k+1=5$-year period) has been analyzed above, in Fig. 5

The phase-break corresponds to the 95% confidence intervals of the corresponding point estimates

As a sensitivity analysis of the results obtained, we estimated Models 2, 3, and 4 for different lags (k=2, 3, 5, 6). Our results, reported in Table 2, are consistent with those presented in Fig. 4, net of some variability inherent to this type of analysis. Most importantly, the slight differences that occasionally emerge do not affect the general picture: a significant slowdown (or disappearance) of the formerly existing convergence process occurred in the observed period.

Discussion

In Italy, survival disparities between the centre-north and the south and islands emerged only in relatively recent years, around the 1990s. These disparities are today even wider in terms of healthy life expectancy, with large differences (up to 4 years between the best and the worst regions), and with a tendency to increase over time (Caselli et al., 2021). In recent years, measures of mortality attributable to the healthcare system (amenable mortality) have been introduced to assess its effectiveness. These measures account for the part of mortality due to deaths that should not occur in the presence of effective and timely care, and for which there are proven, effective therapeutic interventions. In Italy, two sets of regions are associated with significantly lower- and higher-than-average amenable mortality values. Most central-northern regions, except for Valle d'Aosta and

Piemonte, belong to the former group, while Lazio, Abruzzo, Campania, Calabria, and Sicilia belong to the latter (Fantini et al., 2012, 2019). Globally, indications are that the healthcare system does not function equally well all over the country.

A key result of our study is the identification of a halt in the convergence process of life expectancy across Italian regions. This halt occurred in the years between the early 1990s and the early 2000s, in correspondence with the beginning of the regionalization process of the Italian healthcare system, a pattern analogous to that observed in Spain by Montero-Granados et al. in 2007. While we do not know why it happened, our analysis indicates that the phenomenon exists, which, we believe, merits further research.

The halt in the convergence process seems to have occurred earlier for females (around 1993) then for males (around 2001). However, the confidence intervals that we estimated are large, especially for women, and the distance between these two breaks may be smaller than the 8 years that a simple difference would suggest. Simultaneity is not to be expected, however: the chronology of mortality evolution and changes typically differs between males and females. In the 1970s, and for some years thereafter, the convergence process was faster for males than for females: this had been noted before (Casacchia & Natale, 2011; Caselli et al., 2021) and our analyses confirm it (Figs. 4, 5, 6). An important factor in this process seems to have been the cardiovascular revolution, i.e., the rapid progress in medical procedures aimed at combating cardiovascular diseases (beta-blockers, coronary bypasses, stents, etc.) that began in the 1970s, and that benefited mainly males (Bots et al., 2017). Besides, smoking started to decline among males in those years, while it increased among females (Janssen, 2020). Hence, the regional convergence process in male life expectancy may have responded more slowly to the transformations in the healthcare system because of the simultaneous action of other processes (cardiovascular revolution, reduction in smoking prevalence, etc.). Another possible interpretation is that women are more reactive to changes in the healthcare system, because they tend to interact with it more than men do (Benyamini et al., 2000; Pinkhasov et al., 2010), in part because, while living longer, they suffer from poorer health (the male–female health-survival paradox; Alberts et al., 2014; Crimmins et al., 2011; Oksuzyan et al., 2008).

Given our initial hypotheses, we expected to identify a structural breakpoint following the implementation of the 2007 financial recovery plans (Fig. 5), which would have been consistent with the findings of several studies on this topic, as discussed in the introduction. Actually, after 2010, our series do display values that are slightly above the period mean (Fig. 5), but the gap is not large enough for our technique to mark it as significant. This may be due to the brevity of the time series after 2007, possibly insufficient to meet the demanding requirements of the econometric technique adopted here, especially when these effects are small. Unfortunately, the few years that we have past 2019 reflect primarily the death crisis of the COVID-19 pandemic, and are therefore unfit for the fine analyses that we are attempting here.

Our results do not contradict, but rather complement those identified by the literature on this topic. Others before us noted a limited progression, if not a stagnation, in the regional survival convergence process (e.g., Caselli et al., 2021; Cavalieri & Ferrante, 2020), but they did not decompose their series into shorter and, we submit, qualitatively distinct subperiods. In our interpretation, what these authors find is an average between at least two distinct phases, where initial convergence (until the nineties) was later replaced by stagnation.

To conclude this work, we can speculate about the possible causal mechanisms, to be explored in future studies, that might connect the regionalization of the healthcare system with the halt in life expectancy convergence in Italy. Vallin and Meslé (2004) theorized that the progression of life expectancy tends to follow alternating phases of convergence and divergence. Specifically, a sudden medical breakthrough, such as the discovery of antibiotics or beta-blockers, should initially lead to divergence, because only a few countries (the forerunners) will promptly adopt it and benefit from its advantages. However, this initial phase of divergence should eventually give way to convergence, as more countries adopt the innovation. While this theoretical framework was initially developed to explain international phases of life expectancy divergence and convergence, it can also be applied to subnational patterns, especially in countries whose healthcare systems have undergone regionalization, as in several parts of Europe. Within this framework, the slowdown in regional convergence in Italy may be attributable to the higher readiness of north-central regions to adopt new medical and sanitary innovations. The regionalization of the healthcare system may thus have enabled north-central regions to fully leverage their greater openness to medical and sanitary progression.

Alternatively, it can be hypothesized that regionalization may have reduced financial transfers to the poorer regions in the south. An analysis of how the central government allocates resources to regional healthcare systems suggests that regionalization coincided with a shift in resource allocation that favored north-central regions (see Ciocci & Spagnolo, 2020). Until recently, this redistribution was based primarily on population size and, to a lesser extent, population aging. However, lower average income, poorer health conditions, and lower life expectancy in south regions were not considered. From the outset, the regionalization of the healthcare system seems to have been implemented without a system of financial transfers aimed at promoting convergence.

It is possible that both factors—the greater openness of north-central regions to innovation and the unfavorable redistribution of economic resources to the south—have contributed to the end of regional convergence in Italy. For now, we can state with reasonable certainty that a halt in the process of regional convergence has indeed occurred, which raises doubts as to the effectiveness of the national health system.

Author contributions

Federico Benassi provided the original idea for the study, and later contributed to refining the manuscript in various parts. Gianni Carboni, who collected and prepared the data, authored the first draft. Giambattista Salinari added methodological expertise, which contributed to the manuscript's refinement. Gustavo De Santis offered comprehensive critical feedback, and helped restructure and reshape the original draft. All authors extensively interacted during the various stages of preparation of the paper, and eventually agreed on its final version.

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Availability of data and materials

The data used in the paper can be freely download from the ISTAT web site (www.istat.it).

Declarations

Ethics approval and consent to participate

Not applicable.

Competing interests
The authors have no conflict of interest to declare.

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