



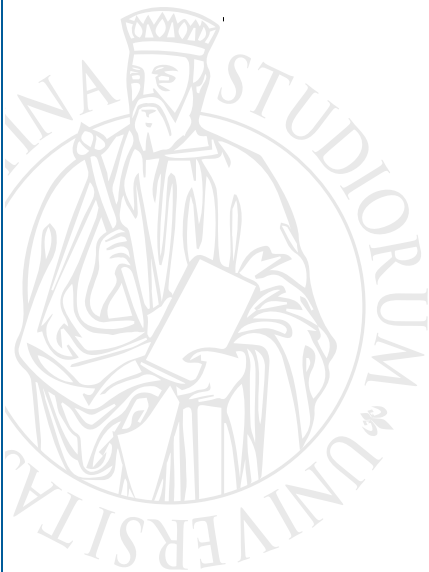
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Mortality evolution in Italy: whatever happened to regional convergence?

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Mortality evolution in Italy: whatever happened to regional convergence?

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Abstract

In several European countries, life expectancy has progressed little in the past two decades. In this paper, focused on Italy, we investigate whether “austerity” and health regionalization may have contributed to this outcome. We show that the succession of reforms to the Italian health system introduced since the 1990s closely corresponds to discontinuities in the evolution of regional life expectancies, halting or reversing their previous trend towards convergence. This holds for both sigma and beta convergence and for both genders, albeit earlier for females.

Keywords: mortality; regional convergence; beta convergence; sigma convergence; public health; health regionalization

1. Introduction

Since the 2000s, progress in life expectancy has markedly slowed down in several European countries. In the UK, for instance, after about 70 years of progress, life expectancy stagnated after the Great Recession of 2008 (Murphy et al. 2019) and it even declined in some years (Newton et al. 2016). Slower survival progress has been noted also in other countries, such as Switzerland, the Netherlands, Belgium, France, and Germany (Franklin et al. 2017, Murphy et al. 2019, Raleigh 2019). Several possible explanations can be advanced for this slowdown, not all of them fully explored: for instance, population ageing, new flu variants, tempo effects,¹ cohort effects,² and migration flows. We submit that this list should be lengthened to include “austerity”, i.e. the set of budget-control policies put in place after the 2008 economic crisis to limit, and possibly halt and reverse, public deficit. Not surprisingly, such policies affect the public resources devoted to health care. This, at least, is what happened in Italy, whose austerity policy began well before the onset of the Great Recession, following the financial crisis of the early 1990s. In those years, Italy was on the verge of defaulting on its large public debt³ and the pressure of national and international investors was huge. In 1992, a “technical” government was charged with the responsibility of a series of draconian measures to bring national finances back under control. Among other things, the budget of the national healthcare system was limited,⁴ and control was partly decentralized to local authorities, the 20 administrative regions in which the national territory is divided.⁵

In 2001, a hefty constitutional reform reinforced the tendency towards “regionalization”, and health fell under the joint responsibility of the central state (charged of specifying the so-called *Minimum Levels of Care*), and regions, deciding on all practical matters concerning the regulation and organisation of health services, including the management of financial resources.⁶

These two reforms ended up by making the Italian healthcare system more fragile and fragmented, and in this paper we hypothesize that this affected the slower survival progress of subsequent years (e.g. Egidi and Demuru 2018). Between 1970 and 2008, life expectancy in Italy had increased along a more or less linear trend, with a slope of about 0.27, that is by almost three years of life every ten

¹ According to Luy et al. 2020 (p. 98), “Tempo Effects (TE) emerge in death rates as soon as mortality is changing during the observation period. For example, when improvements in health and living conditions lead to a reduction in mortality, a certain number of deaths – which would have occurred under unchanged mortality conditions – are postponed to a later period. Such a postponement consequentially deflates the numerator of death rates by the number of avoided deaths, while the denominator is inflated by the same number of saved lives. Albeit these numbers are identical in absolute terms, they differ notably in relative terms. The number of deaths decreases by a much larger proportion than the number of the population at risk in the denominator increases. In this way, TE magnify the effect of the shifted number of deaths in the death rate. The larger the changes in mortality, the larger the magnification effect”. Obviously, the same occurs with variations of a positive sign.

² This holds especially for the UK. Compared to earlier and later generations, the UK birth cohorts of the 1930s (sometimes called “golden cohort”) have had markedly high levels of survival, so much as to affect overall mortality metrics. With the progressive reduction of the survivors of this cohort, this effect has gradually diminished, and has virtually vanished nowadays.

³ In September 2002, the Italian Lira, together with the British Pound, was forced out of the Exchange Rate Mechanism - ERM.

⁴ The Italian National Health Service was created in 1978 and became fully operative in the 1980s. It was intended to provide universal and free health to the entire population, but costs mounted rapidly so that various forms of co-payment have been progressively introduced over time.

⁵ Italian regions constitute the first level of territorial partition of the Italian state as well as public entities endowed with political and administrative autonomy, as stated by the Constitution of the Italian Republic.

⁶ For an overview of the Italian health care system and its reforms, see Ferrè et al. (2014) and de Belvis et al. (2022)

calendar years. In the post-crisis period (2008-2019), the gain declined substantially, to only two years every ten years (Salinari et al. 2023).

In this paper, we analyse a specific part of this complex story: the evolution of Italian regional life expectancies between 1974 and 2019. We set out to show that the pre-existing tendency towards convergence of these regional life expectancies halted, and in some cases reversed, in the reform years, or shortly after. The decision to use administrative regions as statistical units stems directly from the nature of these reforms, which conferred regions more and more autonomy and decision-making power in the health field.

The topic is not entirely new. In 2011, Casacchia and Natale, in their analysis of the evolution of regional life expectancies in Italy over the period 1974-2008, had already noted convergence, especially among males. However, they did not try to relate this to policies and reforms. Cavalieri and Ferrante (2020) did, and concluded that decentralization had probably favoured homogenization, i.e., convergence of regional life expectancies at birth. As we will show shortly, these conclusions run counter our own findings, and those of Montero-Granados et al. (2007) for Spain, possibly because of their shorter period of investigation (1996 to 2016), or of their different methodology, or both.

We investigate more or less the same variables as Casacchia and Natale (2011), with two innovations. First, we cover a longer period, 1974 to 2019, which permits us to compare the pre-reform phase (before the 1990s) to subsequent years. Although data for even more recent years are also available, we excluded them from the analysis, because the outbreak of the Covid-19 pandemic affected the level and geographical distribution of mortality in Italy so profoundly as to obscure all other possible underlying causes.

Secondly, we apply not only beta and sigma convergence analysis, but also a more refined econometric technique, proposed by Bai and Perron (2003), to identify significant changes (or “structural breaks”) in time trends. In our case, the time trend of interest is that of regional convergence in survival, as measured by life expectancy at birth (e_0).

2. 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 consider e_0 separately by gender. The territorial units of our analysis are administrative regions (or NUTS 2). There are 20 of them, but as Valle d’Aosta and Molise are very small, we merged them with the neighbouring regions with which they were originally united: Piedmont and Abruzzo, respectively. This leaves us with 18 regions, the convergence of whose life expectancies we study in two ways, beta-convergence and sigma-convergence, borrowing a methodology that was first developed in economics (Barro and Sala-i-Martin 1990 and 1992; Barro et al. 1991) and later exported to demography (Hrzic et al. 2021).

The most intuitive approach to the study of convergence is to examine how a dispersion index changes over time: if dispersion decreases, sigma convergence occurs. As an index of dispersion for regional life expectancies at birth which we will generically denote by e_r , we opted for variance, formally:

$$VAR(e_t) = \frac{1}{R-1} \sum_{r=1}^R (e_{t,r} - \bar{e}_t)^2, \quad t = 0, \dots, T$$

for each region r (of which we have $R=18$), and each year t , ranging from 0 (first year of observation, 1974) to $T=45$ (2019).

As the population of Italian regions differs widely (from some 1.5 million in Abruzzo and Molise to some 10 million in Lombardia), following Hrzic et al. (2021), we weighted the dispersion index using regional populations, as follows:

$$WVAR(e_t) = \frac{W'_t}{W''_t - W'_t} \sum_{r=1}^R P_{t,r} (e_{t,r} - \bar{e}_t)^2, \quad t = 0, \dots, T$$

where $P_{t,r}$ is the population of region r at time t , $W'_t = \sum_{r=1}^R P_{t,r}$, and $W''_t = \sum_{r=1}^R P_{t,r}^2$. We computed confidence intervals with non-parametric bootstrapping (Mills and Zandvakili 1997).⁷

Sigma convergence gives an idea of the evolution of survival variability across regions, but it misses some potentially important elements. For instance, in the extreme case of ranking reversal, with top regions becoming last and vice versa, dispersion could remain (almost) the same, falsely signalling that (almost) nothing has changed. To overcome this potential limitation, following Barro et al. (1991), we also analyze the so-called beta convergence, checking whether regions with low initial values experience 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 and Theurl 2011).

Let us define the following two quantities for a generic time series y_0, \dots, 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.

The existence of beta convergence can be assessed by looking at the β parameter in the following model:

$$\frac{1}{T+1} \Delta_T e_{T,r} = \alpha + \beta e_{0,r} + \varepsilon_r, \quad r = 1, \dots, R \quad (1)$$

where $\frac{1}{T+1} \Delta_T e_{T,r}$ is the average annual increase in life expectancy between 0 and T for region 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 . We can use the value of life expectancy directly, and not its logarithm as economists frequently do because, unlike economic variables (such as the GDP), life expectancy tends to follow a linear rather than an exponential trend (Oeppen and Vaupel 2002, White 2002, Lee 2019). More precisely, given the heterogeneity of Italian regions in terms of population size, we apply an extension of Model 1 and

⁷ To check the robustness of our analyses, we repeated the same calculations with Theil's index instead of the variance. Results, available upon request, are almost identical and are not reported here. Note that e_0 is a regional average: our analysis ignores internal (i.e. intra-regional) variability.

estimate a population-weighted regression (WLS), where, as earlier, the weights $w_{t,r}$ represent the population size of each region r in year t .

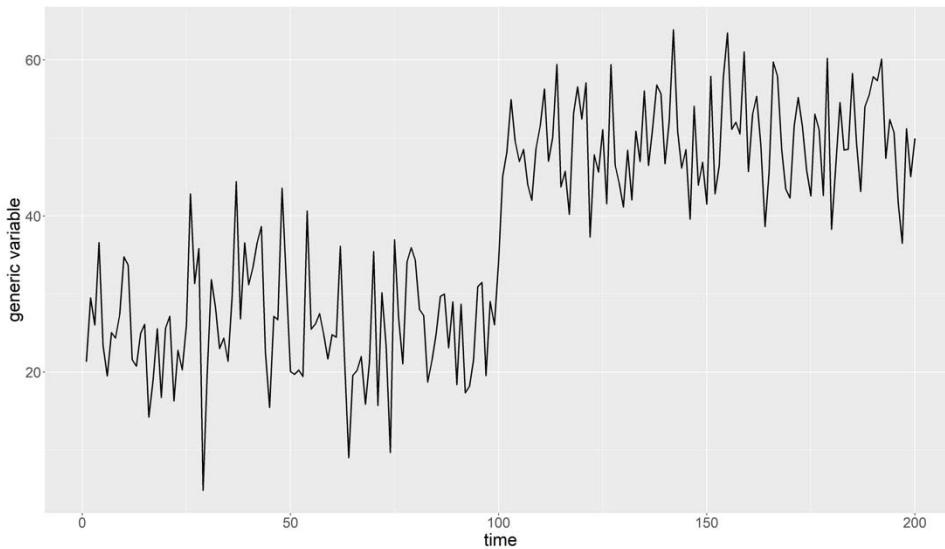
We also created a second time series of beta coefficients, applying the weighted version of Model 1 to shorter (five-year) time intervals, to assess what happened within each sub period:

$$\frac{1}{k+1} \Delta_k e_{i,r} = \alpha + \beta_i e_{i-k,r} + \varepsilon_r, \quad r = 1, \dots, R; \quad i = k, \dots, T \quad (2)$$

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

We want to test whether our $\boldsymbol{\beta}$ parameters behave like a shift model, initially oscillating around a certain mean c' and then, after a jump (or slump), around a different mean c'' . Graphically, a generic shift model behaves as in Figure 1, and can be conceived as the juxtaposition of two different stationary series around their (different) means.

Figure 1 - Hypothetical trend of a shift model on simulated data



Source: authors' simulation

In our case, the shift signals a change in the intensity of the convergence process. If the sign of the average also changes, this indicates 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:

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

The former equation represents the dynamics of the process up to \hat{j} , and the latter the dynamics after it, where \hat{j} is a generic (possible) breakpoint. Note that ε_j and η_j need not be white-noise processes,

as it is usually assumed in standard OLS models: for instance, they could follow an ARMA process, with Auto Regressive and Moving Average properties.

Bai and Perron's technique identifies structural breaks by systematically analysing all possibilities. After choosing a first potential breakpoint \hat{j} , we estimate the following model:

$$\beta_j = \gamma_0 + \gamma_1 B_j + \omega_j, \quad j = k, \dots, T \quad (4)$$

where

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

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 based on \hat{j} .

Next, we choose another potential breakpoint, say \hat{j}' , then a third, \hat{j}'' , etc. and we proceed 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 can now verify whether the estimated β 's behave like a shift model (Model 3) 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 seems to be stationary. To assess non-stationarity, we used the PP test (Phillips and 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 first identifying the ideal one-break partition, then the ideal two-break partitions, etc. (Bai and Perron 2003). We performed this procedure, using the R library *strucchange* (Zeileis et al. 2002, 2003, Zeileis 2006). We also performed a sensitivity analysis to make sure that the choice of different values for k would not bias our results (see Table 2).

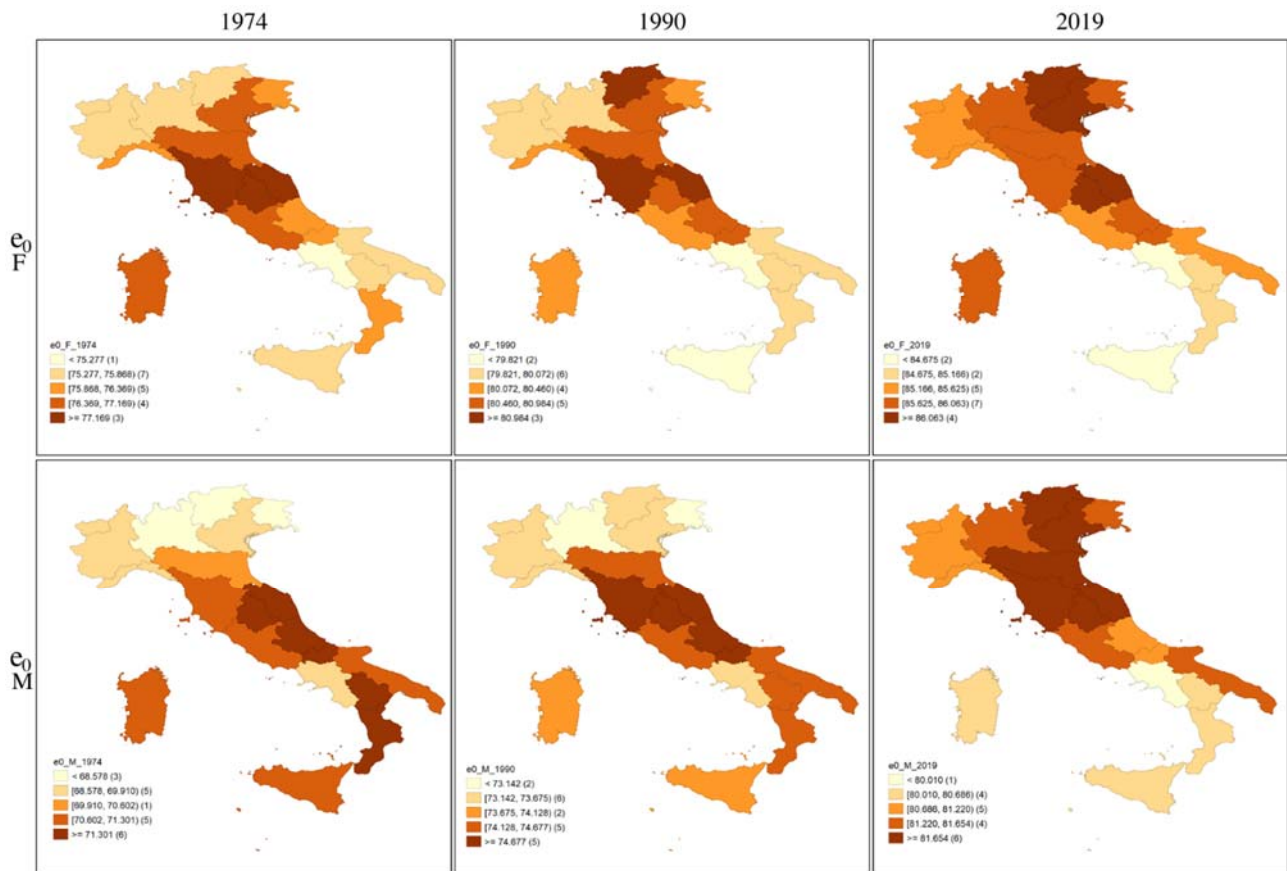
3. Results

In 1979 the average (weighted) life expectancy at birth in Italy was 75.9 and 69.7 years for females and males, respectively, and in 2019 it went up to 85.3 and 81, an increase of 12.4% and 16.7%, respectively.

Currently, life expectancy is lower in southern Italy, mainly because of a greater incidence of cardiovascular diseases (Petrelli et al. 2019) and higher smoking prevalence (Gallus et al. 2011). In Italy, as elsewhere, low social economic status goes along with higher death risks (Frova et al. 2021), but social disparities in mortality are more pronounced in the south (Petrelli et al. 2019).

In 1974, however, the situation was somewhat different, with some southern regions enjoying higher survival than some northern ones. Figure 2 shows the spatial distribution of e_0 across Italian regions, separately by gender, for three years, 1974, 2019, the first and last year of our series, and 1990, when reforms in the Italian health system started to be introduced.

Figure 2 - Spatial distribution of e_0 by gender (F =females; M =males), in selected years.



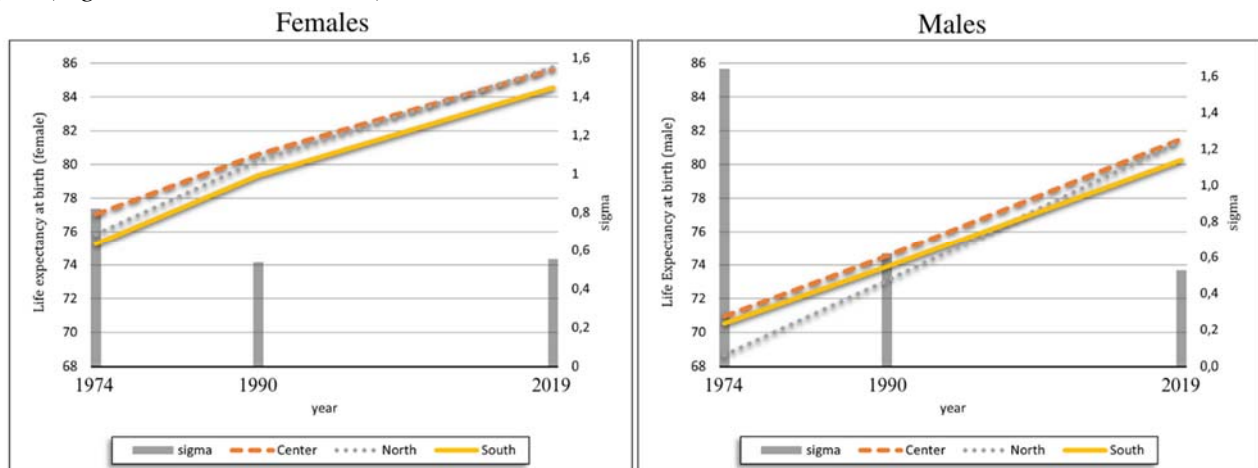
Note: Classes defined using natural breaks method.

Source: Authors' calculations on ISTAT data (dati.istat.it)

The ranking of regional life expectancies changed between 1974 and 2019, and especially around 1990. Before then, that is before the start of the reforms, convergence prevailed; but this process stopped or even reversed thereafter.

Figure 3 shows the evolution of life expectancy in the three traditional macro-territorial areas of Italy (North, Centre, South) between 1974 and 2019, along with the weighted variance or 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) lagged behind in the South, and the distance from the rest of Italy increased for males.

Figure 3 - Weighted averages of e_0 by macro-area (left-hand vertical scale), and sigma convergence of e_0 (right-hand vertical scale)

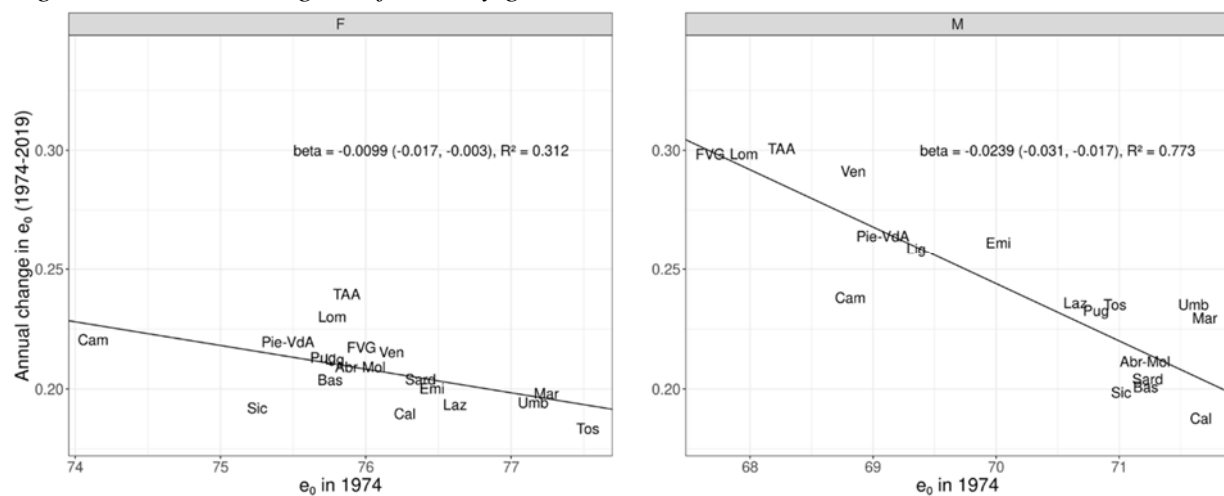


Source: authors' calculations on ISTAT data (dati.istat.it)

To sum up, net of some differences by gender, the general message conveyed by Figures 2 and 3 is that sigma convergence has slowed down since the 1990s, 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, where we also note that convergence has been stronger for males than for females.

Figure 4 - Beta convergence for e_0 , by gender

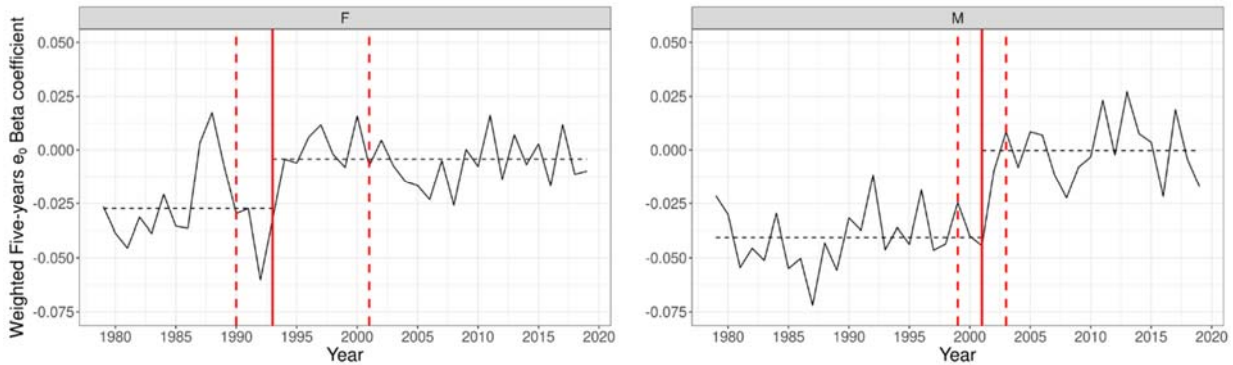


Note: Abr-Mol = Abruzzo-Molise; Bas = Basilicata; Cal = Calabria; Cam = Campania; Emi = Emilia-Romagna; FVG = Friuli-Venezia Giulia; Laz = Lazio; Lig = Liguria; Lom = Lombardia; Mar = Marche; Pie-VdA = Piemonte-Valle d'Aosta; Pug = Puglia; Sard = Sardegna; Sic = Sicilia; Tos = Toscana; TAA = Trentino Alto Adige / Südtirol; Umb = Umbria; Ven = Veneto

Source: authors' calculations on ISTAT data (dati.istat.it).

Figure 5 displays the evolution of the beta-convergence series estimated over five-year time intervals (Model 2), with their corresponding breakpoints (Models 3 and 4, with $k=4$). We estimate a break for females in 1993 (1990–2001 95% CI), and a later break for males, in 2001 (1999–2003 95% CI). These two breakpoints coincide with the two major health reforms mentioned earlier, although the reasons for the earlier occurrence of the break for females are still unclear.

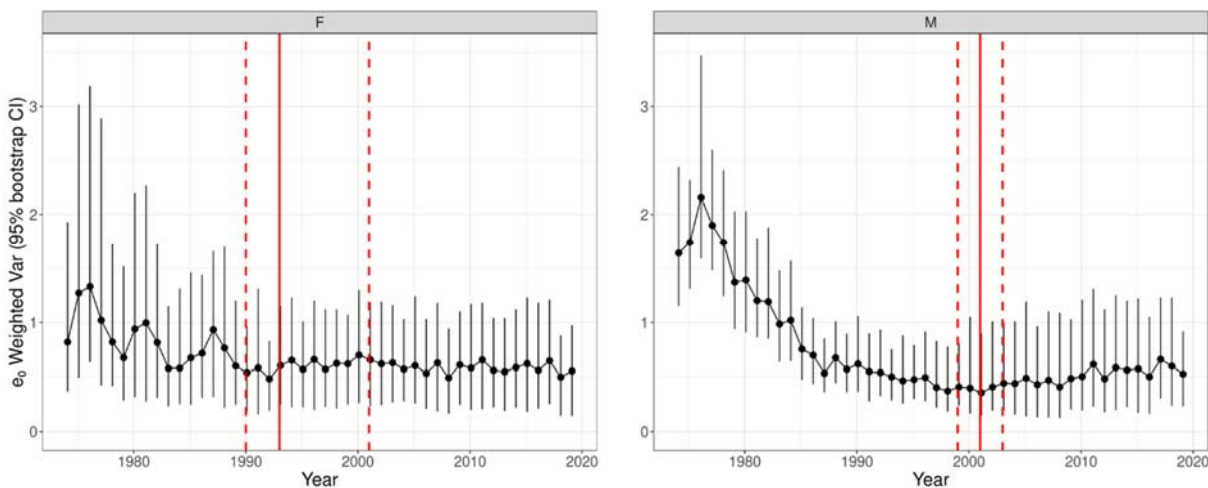
Figure 5 - Time series of the five-year beta convergence for e_0 , by gender. Vertical lines represent our estimates of the structural breaks (continuous lines: point estimates; broken lines: 95% confidence intervals)



Source: authors' calculations on ISTAT data (dati.istat.it).

Figure 6 displays the sigma series together with the estimated breakpoints of Figure 5.⁸ Consistently with expectations, these breakpoints separate an earlier phase of convergence (negative values of the beta coefficients) from a subsequent phase of relative stability.

Figure 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).

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. The assumptions we made in the quest for break points (Model 3) seem consistent with our data, and the beta series do behave like a shift model (Figure 1). Table 1 shows the results of the PP-tests for each variable explored.

⁸ We cannot apply here the same breakpoint-identification technique that we used for the beta series, because the sigma series do not show any clear trend.

Table 1 - Results of PP-tests launched on the residuals of model 4

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 truncation lag parameter is the number of past observations considered in the stationarity-test for each current period.

Source: authors' calculations on ISTAT data (dati.istat.it).

As a sensitivity analysis of the results obtained, we repeated 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 Figure 5, net of some variability inherent to this type of analysis. Most importantly, the slight differences that occasionally emerge do not affect the general picture of a significant change in the convergence process in conjunction with one of the two major health reforms of the Italian system discussed earlier (1992 and 2001).

Table 2 - Structural breaks estimated in beta convergence for different values of k , with the corresponding confidence intervals (sensitivity analysis).

k (Lag parameter)	Females		Males	
	Breakpoint	95% CI	breakpoint	95% CI
2	-		1999	1996-2004
3	1985	1979-1992	2001	1999-2004
4	1993	1990-2001	2001	1999-2003
5	1995	1992-2001	2002	2001-2003
6	1995	1993-2002	2003	2002-2004

Note: “-“ indicates that no significant break was found. We remind our readers that $k+1$ (in years) is the length of the interval under study in that analysis. The case with $k=4$ ($k+1=5$) has been analysed above, in Figure 5.

Source: authors' calculations on ISTAT data (dati.istat.it)

4. Conclusion and discussion

In Italy, the process of convergence of regional life expectancies halted in the final decade of last century. Shortly later, in 2008, in correspondence with the onset of the economic crisis, the progress of survival slowed down markedly (Salinari et al. 2023).

Policy reforms and economic crises are not the sole potential causes of these evolutions. For instance, it could be argued that as Italy is one of the world countries where life expectancy is highest (currently, more than 80 years for men and slightly less than 85 years for women), this deceleration may be the inescapable consequence of approaching the limits of survival. However, whether these limits exist is still a matter of debate and in all cases they have not been identified yet (Linh et al 2023). Besides, there are countries where life expectancy is even higher than it is in Italy, Japan for example, and where no marked slowdown in the progress of life expectancy has been detected.

Our hypothesis is that the reforms of the Italian health system that took place starting in the 1990s played a role in this process. The purpose of these reforms was to limit the waste of public resources, to make regional governments more accountable, and to improve the efficiency of the public health system. In their practical application, however, these reforms created 20 different regional health

systems in Italy, which, we believe, had non-trivial health consequences. To test this hypothesis, we investigated the evolution of convergence in regional life expectancies. The (relatively simple) econometric techniques that we used led us to identify the existence of significant discontinuities (or “structural breaks”) in the convergence process, and these discontinuities emerged precisely in this period of reforms. In short, our hypothesis is not rejected by empirical data.

On the other hand, our hypothesis is not, and cannot possibly be, *proved* by our analysis, which is not intended to reveal the existence of causal links, in this case between systemic reforms and the evolution of regional survival (convergence or divergence, and at what speed). While this crucial link awaits more in-depth studies, our results merit consideration, we believe, for at least two reasons. First, they reveal, and temporally locate, a change of pace in regional survival that had gone unnoticed until now. Incidentally, as Felice et al. (2016) show, the dynamic of convergence (or divergence) in life expectancy seems to anticipate economic fluctuations.

Secondly, we offer a key to interpretation (“regionalizing” a national health system may jeopardize the health homogeneity of a country) that policy-makers may want to take into consideration when deciding on the future of national health systems.

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