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Seasonal components of life expectancy change in Belgium, 2000–2019

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Abstract

BACKGROUND

Although previous studies have examined both short- and long-term changes in mortality seasonality, no research has assessed how shifts in season-specific mortality translate into annual changes in life expectancy while simultaneously accounting for variation by age and cause of death.

OBJECTIVE

The study aims to (1) quantify how short-term changes in mortality by age, cause, and season influence year-to-year changes in life expectancy, and (2) determine the contribution of shifts in seasonal mortality by age and cause of death to the overall increase in life expectancy in Belgium in 2000–2019.

METHODS

Using Belgian population register data, I aggregated deaths by year and quarter, age, sex, and cause of death. I applied an extended decomposition method based on Arriaga's approach, which decomposes differences in life expectancy by age, cause, and quarter of the year.

RESULTS

Year-to-year fluctuations in life expectancy were driven mainly by mortality changes in the first quarter of the year, especially at older ages, from cardiovascular, respiratory, and other diseases with high seasonal variation in mortality. Over 2000–2019, mortality reductions in all seasons contributed to the rise in life expectancy, but declines in the first quarter made slightly larger contributions than those in other quarters.

CONTRIBUTION

I established the link between changes in age-, cause-, and season-specific mortality and changes in life expectancy by adapting the conventional life expectancy decomposition method to account for season-specific mortality differences.

KEYWORDS: mortality, cause-of-death analysis, life expectancy decomposition, mortality seasonality, short-term mortality fluctuations

Introduction

Historically, researchers have studied mortality seasonality independently of annual mortality. A few recent studies, however, examine how season-specific mortality affects annual mortality indicators, particularly life expectancy. One study in Italy finds that eliminating seasonal mortality peaks would reduce regional inequality in life expectancy at birth (e_0) by 11.2%, including a 7.5% reduction from eliminating excess winter mortality alone (Marinetti et al. 2025b). Another study shows that the long-term decline in mortality seasonality accounts for 40% of the increase in remaining life expectancy at age 60 over 150 years in Sweden (Ledberg 2020). A third study evaluates the influence of seasonality on life expectancy across 20 European countries from 2000 to 2019 by using a counterfactual scenario without seasonal excess mortality – replacing seasonal death counts with death counts from the lowest-mortality weeks. On average, e_0 would be 1.14 years higher for men and 0.8 years higher for women without season-specific episodes of excess mortality in Europe, although the results vary substantially across countries and periods (Marinetti et al. 2025a). However, to my knowledge, no study has examined how differences in seasonality contribute to differences in annual life expectancy when stratifying simultaneously by age and cause of death. Beyond this, the question of how changes in season-specific mortality patterns – including short-term fluctuations and long-term shifts – shape changes in annual mortality estimates remains largely unexplored.

Belgium provides an informative case study because it combines low mortality concentrated at older ages (nearly two-thirds of those who die have reached 80 years or

more (Eggerickx, Sanderson, and Vandeschrick 2020)) and significantly higher excess winter mortality than the average across 31 European countries when using the ratio of winter (December–March) to non-winter deaths as the metric (Fowler et al. 2015). Between 2000 and 2019, e_0 for both sexes in Belgium increased by 4 years, from 77.8 to 81.8, yet this growth was uneven: at several points, e_0 stagnated or even declined. For example, e_0 fell by 1.2 months in 2012 for both males and females and again by 3.6 months in 2015 for females (Statbel 2025). These episodes of e_0 decline or stagnation can potentially be explained by seasonal events, such as influenza outbreaks and heat waves.

This study has two objectives. First, I assess how year-to-year changes in age-, cause-, and season-specific mortality influence year-to-year changes in annual e_0. Second, I quantify the contribution of changes in seasonality by age and cause of death to the overall increase in annual e_0 from 2000 to 2019. I introduce a method that is technically similar to standard life expectancy decomposition by age and cause of death but, in addition to age and cause, also includes a seasonal component. I chose 2019 as the last year of observation because the pandemic significantly altered mortality trends from 2020 to 2022 (Natalia et al. 2024), and data beyond 2022 were not yet available at the time of analysis.

Data and Methods

Data

I used the population register data provided by Statbel to obtain death records and population stocks to construct all mortality metrics in this study. First, I aggregated death

records by year and quarter of death, sex, age group (0, 1–4, 5–9, ..., 95+), and cause of death. I defined quarters of the year as 13-week periods: Q1 (week 1–13), Q2 (week 14–26), Q3 (week 27–39), Q4 (week 40–52). I classified causes of death into five broad groups:

- Cardiovascular diseases (ICD-10 codes: I00–I99, G45, K64)
- Neoplasms (ICD-10 codes: C00–D48)
- Respiratory diseases (ICD-10 codes: J00–J98)
- External causes of death (ICD-10 codes: V01–Y89)
- Other natural causes of death (referred to as "Other" throughout the manuscript)

Deaths from unknown causes were redistributed proportionally across all known causes. I also redistributed deaths from ill-defined and unspecified causes (ICD-10 codes: R00–R94, R96–R99) proportionally across all non-external causes.

The denominators for annual mortality rates were calculated as mid-year populations. For the quarter-specific rates, I used mid-year populations divided by four – the number of within-year time units.

Age-standardized mortality rates

For the analysis of season-specific mortality trends, I used age-standardized mortality rates (ASMR). To calculate the ASMR, I used weights from the European Standard Population 2013 (Eurostat 2013).

Life table construction

I constructed abridged life tables using standard demographic methods (Preston, Heuveline, and Guillot 2001). The open age interval was set at 95+.

Decomposition method

In this study, to measure the contribution of changes in age-, cause-, and season-specific mortality to the change in annual life expectancy, I slightly adapted Arriaga's decomposition algorithm (Arriaga 1984, 1989). This decomposition can be conducted in several steps.

First, one can decompose by age the total difference between two values of life expectancy at birth. The contribution of age group *[x;x+n]* to the total difference in life expectancy at birth between two populations is calculated using formula (1):

$$ {}_n\Delta_x = \frac{{}_n\Delta_x^{2-1} + {}_n\Delta_x^{1-2}}{2}, \tag{1} $$

where ${}_n\Delta_x^{2-1}$ and ${}_n\Delta_x^{1-2}$ are computed using formulas (2.1) and (2.2), following the notation given in Preston, Heuveline, and Guillot (2001):

$$ {}_n\Delta_x^{2-1} = \frac{l_x^1}{l_0^1} \cdot \left(\frac{{}_nL_x^2}{l_x^2} - \frac{{}_nL_x^1}{l_x^1} \right) + \frac{T_{x+n}^2}{l_0^1} \cdot \left(\frac{l_x^1}{l_x^2} - \frac{l_{x+n}^1}{l_{x+n}^2} \right); \tag{2.1} $$

$$ {}_n\Delta_x^{1-2} = \frac{l_x^2}{l_0^2} \cdot \left(\frac{{}_nL_x^1}{l_x^1} - \frac{{}_nL_x^2}{l_x^2} \right) + \frac{T_{x+n}^1}{l_0^2} \cdot \left(\frac{l_x^2}{l_x^1} - \frac{l_{x+n}^2}{l_{x+n}^1} \right). \tag{2.2} $$

In these formulas, $l_x^1, l_x^2, {}_nL_x^1, {}_nL_x^2, T_x^1, T_x^2$ are the life table functions, with upper indices 1 and 2 representing populations 1 and 2, where *x* denotes the beginning of the age interval,

and *n* denotes the length of the age interval. For the open age interval, the second term of the right-hand side of formulas (2.1) and (2.2) disappears.

The second step involves decomposing these age-specific contributions ${}_n\Delta_x$ by cause of death and within-year time units (such as weeks, months, or quarters). Although mathematically, the procedure appears to be very similar to the age and cause-of-death decomposition described by Arriaga (1989), there are important semantic differences when incorporating within-year time units.

The contribution of within-year time unit *p*, cause *c*, and age group *[x;x+n)* to the total difference in life expectancy between populations 2 and 1 is calculated using formula (3):

$$ {}_n\Delta_x^{p,c} = {}_n\Delta_x \cdot \frac{\left(\frac{{}_nD_x^{2,p,c}}{{}_nE_x^2}\right) - \left(\frac{{}_nD_x^{1,p,c}}{{}_nE_x^1}\right)}{{}_nm_x^2 - {}_nm_x^1}, \quad (3) $$

where ${}_n\Delta_x$ is the age group contribution from formula (1); ${}_nm_x^2$ and ${}_nm_x^1$ are annual all-cause age-specific mortality rates; ${}_nD_x^{2,p,c}$ and ${}_nD_x^{1,p,c}$ are the number of deaths within the time unit *p* from cause *c*; and ${}_nE_x^2$ and ${}_nE_x^1$ are the annual populations exposed to the risk of death. The upper numerical index indicates the population to which each value refers.

It is important to note that ${}_nD_x^{2,p,c} / {}_nE_x^2$ and ${}_nD_x^{1,p,c} / {}_nE_x^1$ are not mortality rates in the usual sense because the numerators refer to events that occurred within time unit *p*, while the denominators contain annual exposures. The interpretation of the ${}_nD_x^{2,p,c} / {}_nE_x^2$ and ${}_nD_x^{1,p,c} / {}_nE_x^1$ is similar to cause-specific mortality rates – they can be summed across *p* and *c* dimensions to obtain the annual rate.

Finally, the total contribution of within-year time unit *p* to the total difference in life expectancy between two populations is calculated using formula (4):

$$\Delta^p = \sum_{x \in X} \sum_{c \in C} {}_n\Delta_x^{p,c}, \quad (4)$$

where *X* denotes the set of age groups and *C* denotes the set of causes of death. In the same way, one can compute ${}_n\Delta_x^c$ by summing over all values of *p*. These results are identical to those obtained with conventional decomposition by age and cause of death. Summing over all values of *p*, *x*, and *c* yields the total difference in life expectancy between two populations.

Further in the document, for simplicity, I denote ${}_n\Delta_x^{p,c}$ as $\Delta_{x,p,c}$, where *x* represents the age group, *p* the quarter of the year, and *c* the cause of death.

The proposed method has several limitations. First, each within-year time unit *p* must be of equal length so that the population has the same amount of time being exposed to the risk of death in each unit. Second, the method cannot be applied in some rare cases where age-specific mortality rates are identical in both comparator populations, which may occur for very small populations, as it will make the denominator of formula (3) equal to zero. Third, rapid population growth or decline within a calendar year can affect the number of deaths observed in the time unit *p*, meaning that variations in observed deaths may reflect changes in population size rather than changes in mortality levels. In such cases, an adjustment should be designed. In this study, population size changes slowly within calendar years, allowing the assumption that changes in the number of deaths reflect only changes in mortality levels.

Confidence intervals

To account for uncertainty, I used 95% confidence intervals (CI95). To compute these intervals, I assume that the observed mortality level results from a stochastic process in which deaths follow a Poisson distribution:

$$D \sim Pois(\lambda) \, .$$

I then used Monte Carlo simulations with the λ parameter equal to the observed number of deaths $D_{t,x,p,c}^{observed}$ for each combination of year *t*, quarter of the year *p*, age *x*, and cause *c*. The CI95 values were obtained from 1,000 simulations. This procedure is similar to methods described elsewhere (Andreev and Shkolnikov 2010; Silcocks, Jenner, and Reza 2001).

Software

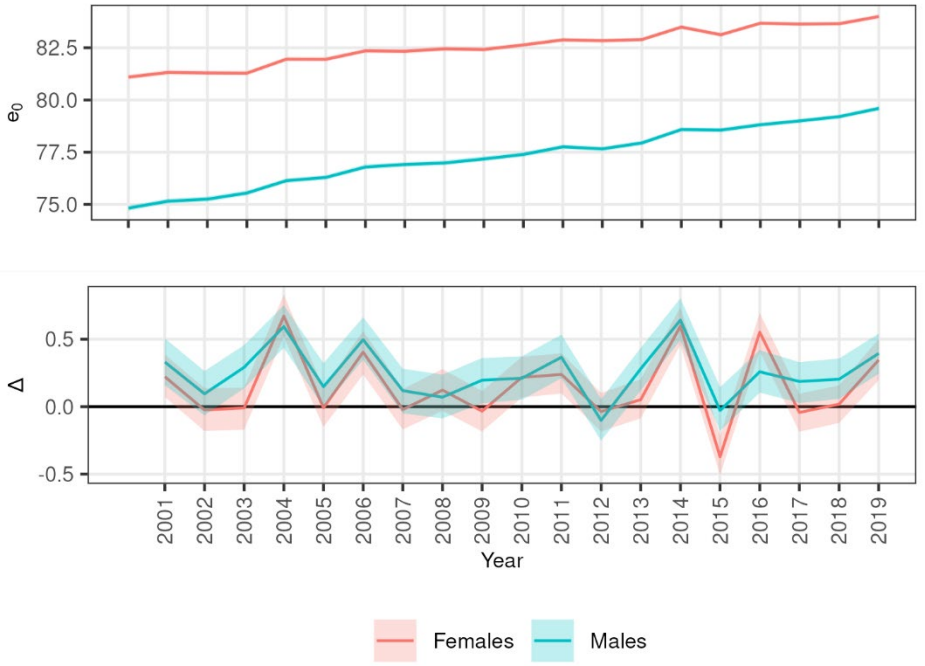
I used R version 4.4.1 for all analyses. I used the "LEdecomp" package (Riffe, Atance, and Lledo 2025) to construct life tables and perform decomposition. For other calculations, I used base R, packages from the "tidyverse" collection (Wickham et al. 2019), and "ggh4x" (van den Brand 2025) package for visualizations.

Results

In Belgium, female e_0 rose from 81.1 years in 2000 to 84.0 years in 2019, while male e_0 increased from 74.8 years to 79.6 years over the same period. However, year-to-year changes in e_0 (Δ) were highly irregular over time (Figure 1). Both sexes experienced

substantial interannual variability in Δ, with alternating periods of faster and slower mortality improvements.

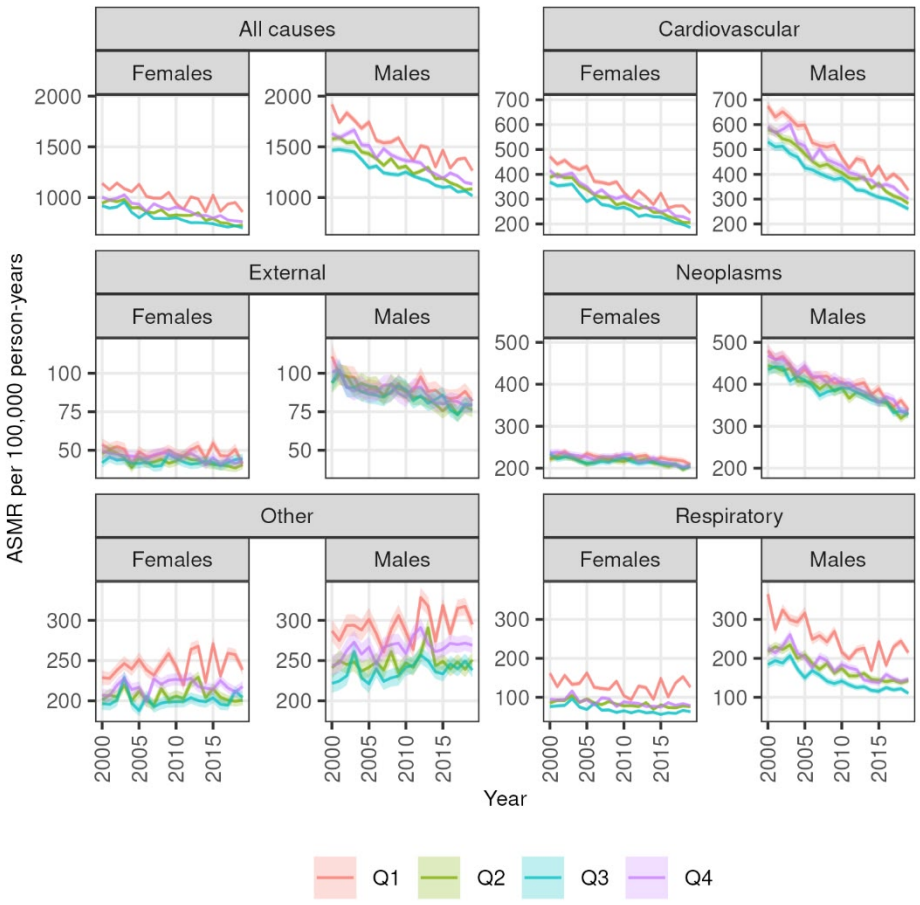
Figure 1: Trends in e_0 and year-to-year changes in e_0 (*Δ*)



Notes: Δ is calculated as e_0(Year) – e_0(Year–1)

Pronounced seasonality was visible in all-cause mortality and in mortality from cardiovascular, respiratory, and "other" causes of death (Figure 2), with ASMRs consistently highest in Q1 and lowest in Q3. Mortality in Q1 showed strong interannual variability. Neoplasms and external causes showed almost no seasonal variation, although both displayed considerable year-to-year fluctuations across all quarters. The Q1–Q3 ASMR difference decreased for cardiovascular mortality among males and females and for respiratory mortality among males, whereas it increased for female mortality from respiratory and "other" causes.

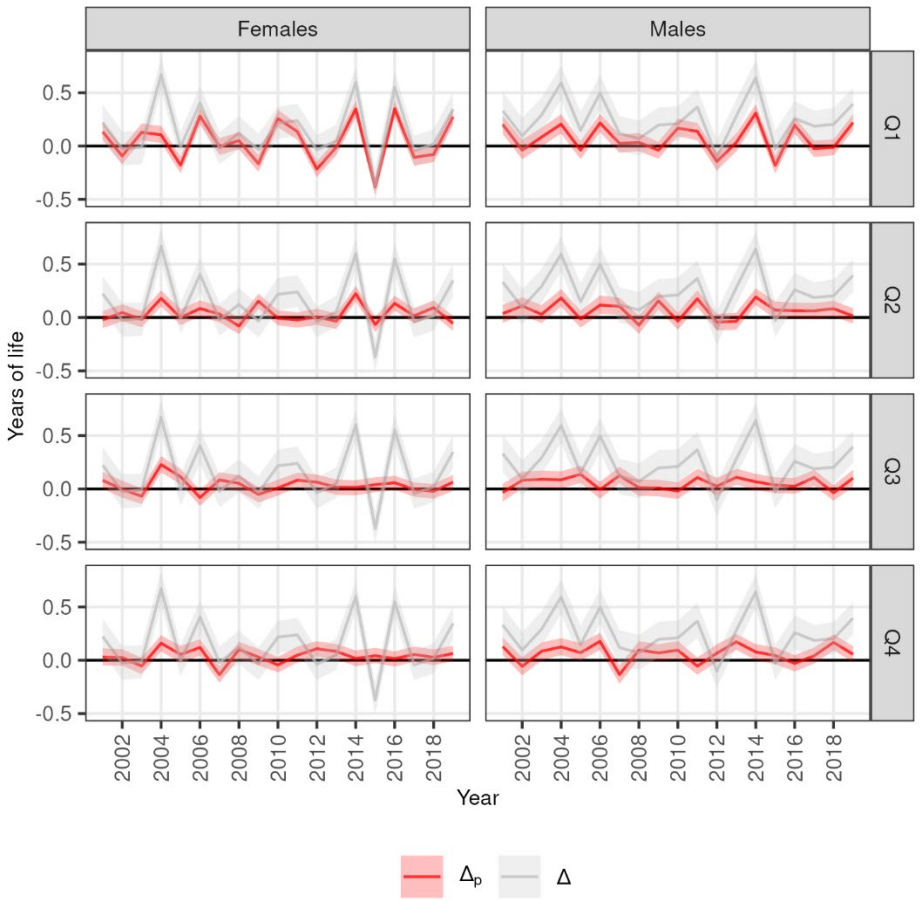
Figure 2: Trends in season-specific ASMR by cause of death



Applying the life expectancy decomposition method to each pair of consecutive years (e.g., 2001 to 2000, 2002 to 2001, etc.) provides insights into how short-term changes in age-, cause-, and season-specific mortality contributed to year-to-year variation in annual e_0. Figure 3 shows that changes in mortality at the beginning of the year – in Q1 and, to a lesser extent, Q2 – were the main drivers of short-term fluctuations in e_0, especially among females in 2010–2019. For example, in 2015, female e_0 declined by 0.37 (0.22–0.51) years relative to 2014; in that year, the increase in Q1 mortality alone accounted for a loss of 0.38 (0.31–0.46) years, while changes in other quarters contributed minimally. A similar

dominance of contributions of mortality changes in Q1 occurred in several other years with significant changes in e_0, including 2010, 2011, 2014, 2016, and 2019. The influence of changes in summer mortality (Q3) on annual e_0 was small throughout the observation period.

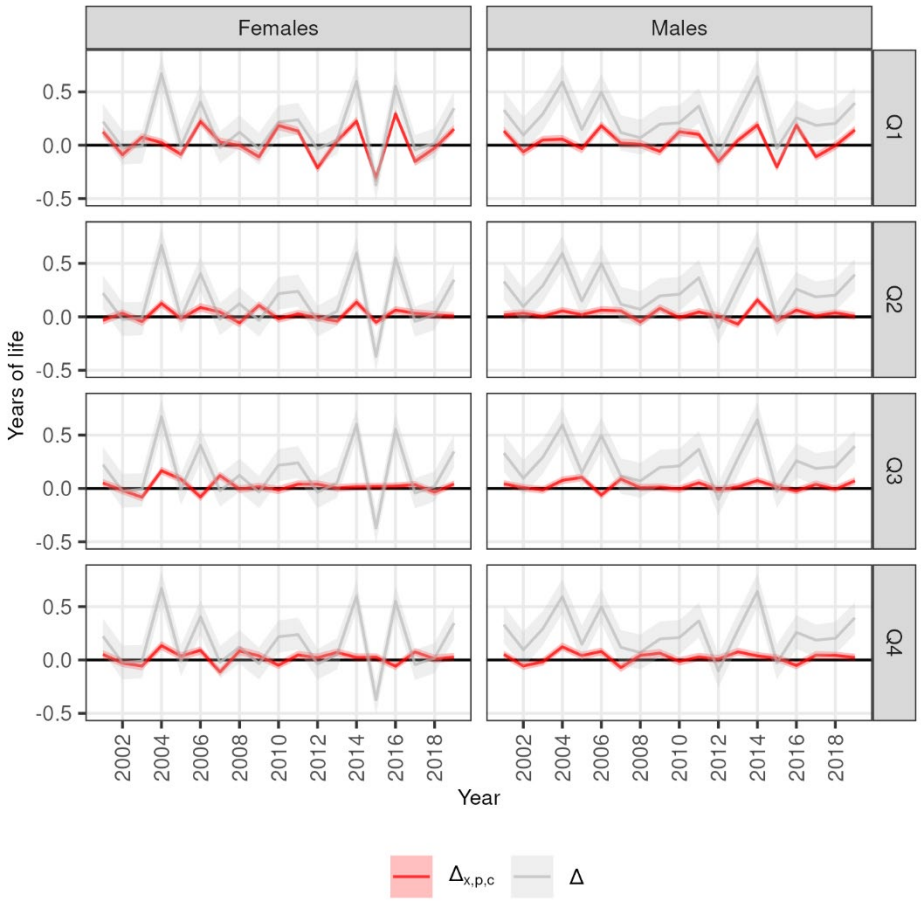
Figure 3: Contribution of year-to-year changes in season-specific mortality (Δ_p) to year-to-year changes in e_0 (Δ)



Notes: Each value is derived by comparing mortality in a given year to mortality in the preceding year (Year–1).

Adding age and cause dimensions allows a more precise identification of which components of mortality drove the year-to-year fluctuations in e_0. The decomposition results show that fluctuations in Q1 mortality among the population aged 60 and over, particularly from cardiovascular, respiratory, and "other" causes, were the main contributors to changes in annual e_0 (Figure 4). In years with the largest differences in e_0 – for example, in 2015 – increases in female mortality in Q1 from cardiovascular, respiratory, and "other" causes among those aged 60+ converted into a loss of 0.30 (0.26–0.34) years in e_0 in relation to 2014, while annual female e_0 declined by 0.37 (0.22–0.51) years. As with the season-only decomposition, one can observe similarities in the time trends of $\Delta_{x,p,c}$ in Q1 and Δ, particularly among females, especially during 2010–2019.

Figure 4: Contribution of year-to-year changes in mortality among the population aged 60 and over from cardiovascular, respiratory, and "other" causes of death ($\Delta_{x,p,c}$) to year-to-year changes in e_0 (Δ)

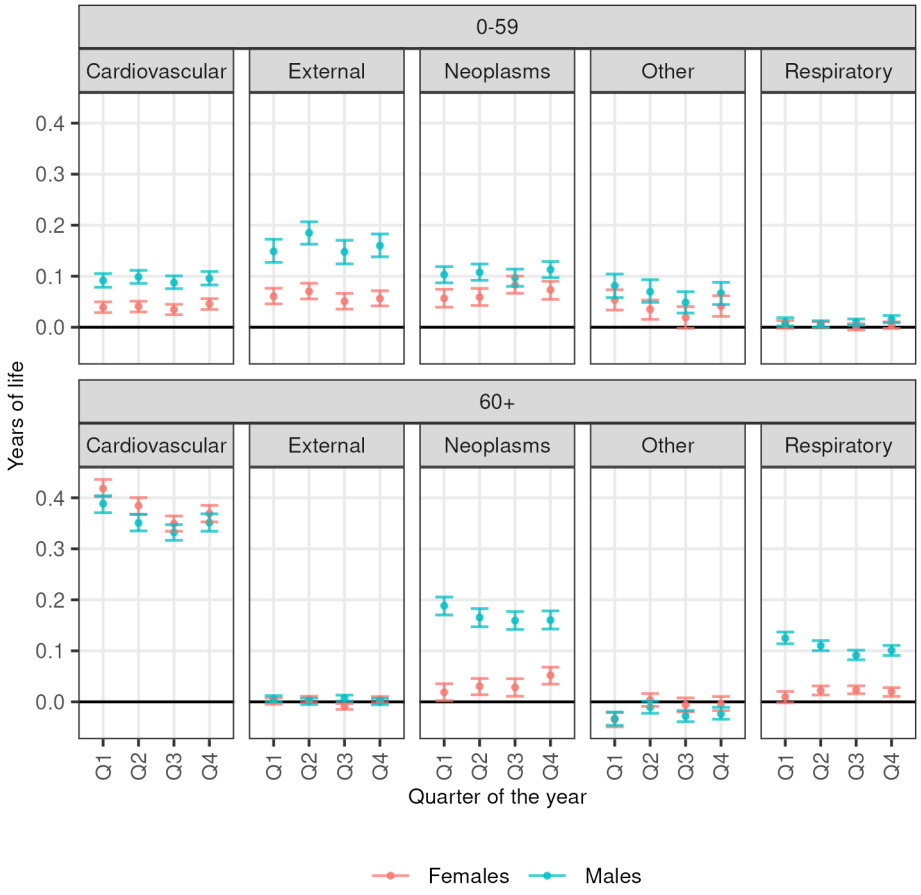


Notes: Each value is derived by comparing mortality in a given year to mortality in the preceding year (Year–1).

To assess how changes in age-, cause-, and season-specific mortality contributed to the overall change in e_0 between 2000 and 2019, I applied a decomposition method to break the total difference in e_0 between 2017–2019 and 2000–2002 into components of interest (Figure 5). The rise in e_0 was driven by all-cause mortality decline across all seasons, with a

slightly larger contribution of non-summer mortality reduction, especially among males. Breaking down season-specific contributions by age and causes of death shows that there are large differences in seasonal contributions for cardiovascular (both sexes) and respiratory (males) diseases at ages 60+. For external causes and neoplasms, where winter mortality does not differ substantially from summer mortality, there were only small differences in seasonal contributions to the change in e_0. The same is evident for the remaining group of "other" causes.

Figure 5: Contribution of changes in age-, cause-, and season-specific mortality ($\Delta_{x,p,c}$) to the overall change in e_0 between 2000–2002 and 2017–2019



Conclusion

There is a vast amount of literature on seasonality in mortality (Fowler et al. 2015; Healy 2003; McKee 1989; Rau 2006; Rousson, Paccaud, and Locatelli 2026; Sakamoto-Momiyama 1978). However, only a few studies have explicitly linked seasonal mortality patterns to annual mortality indices (Ledberg 2020; Marinetti et al. 2025a, 2025b). In this study, I introduced a tool for estimating the exact contribution of age-, cause-, and season-specific mortality differences to the difference in annual life expectancy between two populations, based on an adaptation of a widely used life expectancy decomposition algorithm. The tool may be useful in a broad range of applications. For example, it can be used to assess the extent to which mortality differences within a given season explain mortality disparities between socioeconomic groups defined by education level, employment type, income, or other characteristics.

For Belgium, I show that short-term changes in mortality observed at the beginning of the year, particularly in Q1, from cardiovascular, respiratory, and "other" causes among individuals aged 60 and older are the dominant drivers of year-to-year fluctuations in e_0. These changes in cold-season mortality are likely influenced by fluctuations in influenza activity. Previous studies support this interpretation (Nielsen et al. 2019; Reichert et al. 2004; Shkolnikov et al. 2022). When considering long-term trends, I show that the increase in e_0 in Belgium over the 2000–2019 period was driven by mortality improvements in all seasons, with a slightly more pronounced contribution of reductions in Q1 mortality among males, particularly from cardiovascular and, to a lesser extent, respiratory causes at older

ages. Faster declines in winter mortality from these causes may reflect changes in exposure to cold and heat. This interpretation is consistent with evidence that 2000–2019 was marked by declining cold-related mortality and increasing heat-related mortality worldwide, including Western Europe (Zhao et al. 2021). Nevertheless, winter mortality remained higher than summer mortality throughout the study period: in 2015–2019, the Q1–Q3 all-cause ASMR ratio was 1.29 for females and 1.27 for males.

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