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How Long Will You Be a Widow? Determinants, Trends and Income Gradient in Widowhood Duration^{*}

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Abstract

Understanding the duration of widowhood is essential for individuals and effective widow support policies. However, widowhood duration remains an understudied topic. In this article, we provide a quantitative estimation of the impact of three primary determinants of expected widowhood duration at age 60 within a unified framework: (i) the degree of overlap between male and female mortality distributions; (ii) the age gap between spouses; and (iii) the dependence of spousal mortality. Using French life tables from 1962 to 2070 and simulations based on the Gompertz law and a bivariate Gaussian copula, we assess each determinant's relative influence. Our findings show that ignoring spousal mortalities dependence overestimates widowhood duration by three years, while disregarding the age gap underestimates it by one year. In France, in 2020, expected widowhood duration for females at age 60 is 10.4 years and 5.8 years for males. Despite converging gender life expectancies, our projections suggest that widowhood duration will remain high until 2070: 9.2 years for females and 6.2 years for males. Notably, we identify a negative gradient of widowhood duration along the standard of living distribution.

Key-words: widowhood duration, differential mortality, life tables, Gompertz mortality law, bivariate Gaussian copula.

JEL classification: J11; J12; J14; J18.

Introduction

The death of a spouse is one of the most devastating shocks an individual can face, due not only to the emotional impact of losing a family member but also to the substantial financial risk and far-reaching economic implications for the survivor.¹ Knowing the durations individuals spend in this vulnerable situation is of paramount importance for both policymakers and the individuals themselves. First, widowhood duration holds crucial significance as a key parameter for policymakers, as many countries have implemented public survivor insurance schemes aimed at mitigating the decline in household standard of living following the death of a spouse (OECD, 2018). The duration of survivor benefit provisions has a critical impact on the sustainability of the pension schemes. Moreover, widowhood duration may vary among individuals due to factors such as gender or socioeconomic background, thus potentially increasing the risk of poverty among elderly individuals living alone (Munnell et al., 2020). This risk may be exacerbated if a correlation exists between survivors' life expectancy and their income, as suggested by previous research (Chetty et al., 2016). Policymakers are motivated to address this correlation to both mitigate inequalities and prevent vulnerability. Second, comprehending the expected lifespan of survivors after the death of their spouse plays a vital role in enabling individuals to make informed decisions throughout their life-cycle (De Nardi et al., 2021). It has been observed that both females and males tend to underestimate their likelihood of becoming a widow or widower, primarily because they overestimate the number of years they expect to live alongside their spouse (Holden and Kuo, 1996).² Such underestimations may lead to inadequate savings, whether through public or private means, due to an inappropriate time horizon. This is one contributing factor to the economic challenges faced by widows (Munnell et al., 2020).

Until recently, widowhood duration has received limited attention in the literature, with noteworthy exceptions being the recent work by Compton and Pollak (2021) and pioneering studies by Goldman and Lord (1983) and Myers (1959). These studies, based on U.S. life tables, have examined the historical evolution of widowhood duration, iden-

¹Indeed, at spousal death, at least three mechanisms directly modify household resources: (i) the survivor no longer benefits from the deceased spouse's income; (ii) the survivor loses economies of scale; and (iii) depending on certain criteria, the survivor may receive survivor's benefits.

²This tendency to underestimate remaining life expectancy as a widow is partially attributed to a broader underestimation of one's own life expectancy, as previously noted in the literature (Kutlu-Koc and Kalwij, 2017; Dormont et al., 2018).

tifying variations based on educational attainment and race. While they consider factors such as the overlap between female and male mortality distributions and the average age gap between spouses in the population, they do not delve into the dispersion of age gaps or the dependence of spousal mortality. The literature on couples' joint survival and survivors' life expectancy has attracted considerable interest in the fields of risk and insurance, particularly for annuity computation purposes (Denuit et al., 2001; Frees et al., 1996; Youn and Shemyakin, 2001), as well as in the welfare literature aimed at measuring the effect of spousal death (Leroux and Ponthiere, 2013; Ponthiere, 2016; Strand, 2005).³ These studies do consider the dependence of spousal mortality. However, these two strands of the literature do not focus specifically on widowhood duration as a condition of whether one is either a widow or a widower. In other words, they explicitly address widowhood duration at the macro level but overlook both widowhood duration within the widowed population and disparities across socioeconomic characteristics. Although the literature on widowhood duration has generally failed to fully quantify its determinants, recent interest among demographers has arisen regarding the probability of males outliving females, as in the work of Bergeron-Boucher et al. (2022), which itself built upon a broader research framework measuring the inequality of lifespans between two populations (Vaupel et al., 2021). Both papers highlight that even though, on average, women tend to live longer than men, a notable proportion of men have a significant chance of living longer than women. Despite these valuable insights into the overlap between male and female lifespan distributions, a critical aspect remains unresolved: the duration of outsurvival, which is central to our understanding of widowhood duration. We examine the case of France, characterized by three main elements. First, French women rank high in European life expectancy (OECD, 2023). In 2021, female life expectancy was 85.5 years, compared to 79.3 years for males, while in the OECD, it was 83.0 years for females and 77.6 for males. Second, the gender gap in life expectancy in France is among the highest in OECD countries, excluding former Eastern European countries. Overall, it suggests that widowhood duration is likely to be high. Third, France is also one of the OECD countries with the highest expenditures on survivor's pensions, accounting for 1.7% of gross domestic product in 2017, compared to 1% for the OECD as a whole

³The United States Social Security Administration also provides an interactive software called *Longevity Visualizer* that enables visual-analysis of individual and joint survival probabilities of two persons (Alleva, 2016). Following on the same concept, the American Academy of Actuaries and the Society of Actuaries have jointly developed a similar software available at *Actuaries Longevity Illustrator*.

(OECD, 2018). In 2021, 36% of French retired women benefit from a survivor's pension, which represents nearly half of their pension.

This article estimates the impact of three primary determinants of widowhood duration within a unified framework: (i) the mortality distributions of both females and males, including their degree of overlap; (ii) the age gap between spouses; and (iii) the dependence of spousal mortality. Through simulations based on Gompertz law and a bivariate Gaussian copula, we quantitatively assess the relative influence of each determinant. Our findings show that ignoring spousal mortality dependence overestimates widowhood duration by three years, while disregarding the age gap underestimates it by one year. In France, in 2020, expected widowhood duration for females at age 60 was 10.4 years and 5.8 years for males. Despite converging gender life expectancies, our projections suggest that widowhood duration will remain high until 2070: 9.2 years for females and 6.2 years for males. Notably, we identify a negative gradient of widowhood duration along the standard of living distribution.

This paper directly contributes to several strands of the literature. First, we apply the same methodology as Compton and Pollak (2021), Goldman and Lord (1983) and Myers (1959) to French mortality tables, documenting both the level of widowhood duration and its evolution in France over the 1962–2020 period. Prior research had focused primarily on the United States. Second, we expand on the outsurvival literature by enhancing the outsurvival indicator, initially defined by Vaupel et al. (2021), namely by introducing a temporal dimension to measure widowhood duration. Third, using Gompertz law simulations and a bivariate Gaussian copula, we further contribute to these strands by estimating the influence of widowhood duration determinants within a unified framework, thus facilitating a comparison of their relative impacts. This includes quantifying, for the first time, the significance of both the dispersion of the age gap between spouses and the dependence of spousal mortality when measuring widowhood duration. Fourth, we project our analysis until 2070, going beyond the historical evolution of widowhood duration as described in prior research (Compton and Pollak, 2021). We not only document that widowhood duration has been notably substantial over the past decades but also highlight that it is projected to decrease, albeit to a lesser extent than what might be inferred from the narrowing gap in average life expectancies between females and males. Fifth, we modestly contribute to documenting the correlation between income and life expectancy (Chetty et al., 2016). We show that widowhood duration decreases linearly with standard of living and is more dispersed in the lower part of the income distribution, increasing the risk of very long widowhood and, in turn, poverty. Furthermore, our proposed approach extends beyond the context of widowhood duration and can be applied to other settings involving two correlated survival times. Specifically, as the spouse is the main caregiver in many cases of disability in a couple, the method can be adapted to measure the duration of informal spousal care using disability-free life tables, thus providing insights into long-term care policies. Our approach can also be utilized to assess the period during which couples have children and its evolution over time, using fertility modeling of both the female and male populations.

Measuring Widowhood Duration

Background

Consider a different-sex couple consisting of spouses of the same age x whose respective ages at death do not depend on each other. Let ω represent the maximum lifespan. The variables for deaths, survivals, and life expectancies at age u are denoted as d(u), s(u)and e(u), respectively. Let index f refers to females and m to males. The life expectancy of the wife at age x is calculated as:

$$LE_f(x) = \int_x^\omega \frac{s_f(u).du}{s_f(x)} \tag{1}$$

Our study focuses on a population of couples, where both individuals are alive at age x, so $s_f(x) = s_m(x) = 1$. The probability that a woman survives at age $u \ge x$ is $s_f(u)$, and the probability that her spouse dies at this age is $d_m(u)$. Therefore, she has a $s_f(u).d_m(u)$ probability of becoming a widow at age u. If this event occurs, she has $e_f(u)$ years left to live. The expected widowhood duration (WD) at age x is then derived from the product of these terms, summed over all possible values of u.

$$WD_f(x) = \int_x^{\omega} d_m(u) \cdot s_f(u) \cdot e_f(u) \cdot du$$
(2)

$$= \int_{x}^{\omega} d_{m}(u) \cdot s_{f}(u) \cdot \left[\int_{u}^{\omega} \frac{s_{f}(v)}{s_{f}(u)} dv \right] \cdot du$$
(3)

In order to carry out an integration by parts, we introduce the intermediate notation: $k_f(u) = \int_u^{\omega} s_f(v) dv$, which verifies $k'_f(u) = -s_f(u)$. The previous expression can thus be written as:

$$WD_f(x) = \int_x^{\omega} d_m(u) . k_f(u) . du = -\int_x^{\omega} s'_m(u) . k_f(u) . du,$$
(4)

which we integrate by parts:

$$[-s_m(u).k_f(u)]_x^{\omega} + \int_x^{\omega} s_m(u).k_f'(u).du = -s_m(\omega).k_f(\omega) + s_m(x).k_f(x) - \int_x^{\omega} s_m(u).s_f(u).du.$$

The first term of this expression is null, as $s_m(\omega)$ is null. Since $k_f(x) = \int_x^{\omega} s_f(u) du$ and $s_m(x) = 1$, we have:

$$WD_f(x) = \int_x^{\omega} s_f(u) [1 - s_m(u)] du = LE_f(x) - JS(x),$$
(5)

where JS(x) is the joint survival at age x. In other words, we measure widowhood duration as the female life expectancy minus the joint survival of the spouses. This duration is computed for all women, whether they outlive their husband or not.⁴ When referring to widowhood duration in everyday language, we typically mean widowhood duration conditional on experiencing spousal death, that is, the ratio between unconditional widowhood duration (*UWD*) and the probability that a woman outlives her husband (*proba*). This conditional widowhood duration is then equal to:

$$WD_{f}^{C}(x) = \frac{\int_{x}^{\omega} e_{f}(u).s_{f}(u).d_{m}(u).du}{\int_{x}^{\omega} s_{f}(u).d_{m}(u).du} = \frac{UWD_{f}(x)}{proba_{f}(x)} = \frac{LE_{f}(x) - JS(x)}{1 - \Phi(x)},$$
(6)

where Φ is the outsurvival statistic introduced by Vaupel et al. (2021).

$$\Phi(x) = \int_{x}^{\omega} s_m(u).d_f(u).du,$$
(7)

 Φ measures the probability that an individual from a low life expectancy group will live longer than an individual with a high life expectancy. In this context, males constitute the low life expectancy group and females the high life expectancy group. $\Phi(x)$ then

 $^{^{4}}$ In other words, we include in the calculation the null widowhood duration of women who died before their husbands.

measures the probability that males outlive females at age x and equals one minus the probability that females outlive males, namely widowhood probability.

Determinants of Widowhood Duration

Mortality Distributions Overlap

When considering mortality distributions, it's important to acknowledge that there is an overlap between the mortality patterns of different groups. If all females were to outlive all males, there would be no male widowhood, and widowhood duration would simply be the difference between the average life expectancies of females and males. This common misconception underscores that widowhood duration is influenced by variations around the average age at death and the degree of overlap between the mortality distributions of females and males. For a more detailed framework explaining the impact of mortality overlap on widowhood duration, see online Appendix A.1, where we show that calculating widowhood duration as a simple difference between individual life expectancies of spouses can result in negative counts when the order of deaths is reversed. However, understanding the extent to which overlap affects widowhood duration cannot be divorced from empirical investigation. The degree of overlap depends on the characteristics of mortality distributions, which have undergone substantial changes in the past and will continue to do so in the future. These changes, including processes such as mortality distribution compression and rightward shifts, are well-documented (Bergeron-Boucher et al., 2015).

Age Gap between Spouses

On average, women tend to marry older men (Drefahl, 2010; Goldman and Lord, 1983). For a given woman, marrying an older man increases her probability of widowhood at a young age compared to marrying a same-age partner, as older men have higher death rates than their younger counterparts. This, in turn, may also increase the duration of widowhood because the younger a woman is at her spouse's death, the longer her expected years as a survivor. This insight underscores the role of the age gap between spouses in widowhood. To provide a formal framework for understanding unconditional widowhood duration, we extend the numerator of equation (6) to accommodate bivariate distributions of the age of the wife and the age of the husband:

$$UWD_f(x) = \int_{u=x}^{\omega} \int_{z=a}^{b} \int_{y=A}^{B} e_f(y+u) \cdot s_f(y+u) \cdot d_m(z+u) \cdot dy \cdot dz \cdot du,$$
(8)

where a and b are the youngest and oldest ages of husbands and A and B are the youngest and oldest ages of wives. In the simplified scenario where the age gap between spouses is constant across the population, we provide more details in online Appendix A.2 showing that we introduce an upward bias into the widowhood duration measure if we assume that a fixed age difference of Δ adds Δ years of widowhood to all women. This assumption overlooks the fact that this difference will not prevent some women from passing away before their husbands, while others may survive less than Δ years after their husbands' deaths. Despite the average age gap remaining relatively stable over time, variations within the population may still play a significant role. In light of this, empirical simulations are warranted to assess the influence of the age gap and its dispersion on widowhood.

Dependence of Spousal Mortality

Spouses often exhibit assortative mating patterns according to their socioeconomic characteristics (Gonalons-Pons and Schwartz, 2017). However, age at death is not randomly distributed in the population, with higher socio-economic status (SES) individuals typically experiencing later mortality than those with lower SES. This implies a correlation in ages at death within couples. The shared risks and standards of living during their common life further amplify this phenomenon. Additionally, the well-documented 'broken heart effect' which leads to increased mortality among widowers (Elwert and Christakis, 2008; Parkes et al., 1969; Shor et al., 2012), underscores that the assumption of independence of spousal mortality is not valid. The field of risk insurance has extensively explored the dependence of spousal mortality as a means to measure joint survival and determine survivors' life expectancy, particularly for annuity calculations (Denuit et al., 2001; Frees et al., 1996; Youn and Shemyakin, 2001). We can empirically examine the role of the dependence of spousal mortality within the context of widowhood duration measurement by using bivariate copula functions commonly used in such calculations.

Data and Methods

French Life Tables

We utilize the life tables from the French Institute for Statistical and Economic Studies (INSEE) for both females and males. These tables offer age-specific mortality distributions spanning from 1962 to 2020, and they include projected mortality tables from 2021 to 2070, specifically for metropolitan France (Algava and Blanpain, 2021). Additionally, we incorporate the 2016 life tables categorized by standard of living as computed by (Blanpain, 2018) using French fiscal data. Blanpain (2018) determines the standard of living ventiles for the entire population, pooling together both females and males, while also accounting for differences across various age groups.

Computation using Life Tables

We calculate widowhood duration by adapting Equation (6) to discrete-time, employing mortality tables for randomly paired couples. Widowhood duration measures how long a surviving spouse can expect to live after the death of their current spouse, without considering remarriage. This approach aligns with existing literature, particularly the work of Compton and Pollak (2021). Our choice of the 60-year-old threshold is grounded in two additional justifications. First, the credibility of assuming no divorce before widowhood and no remarriage after widowhood increases as divorce rates dramatically decrease after age 40 (Kennedy and Ruggles, 2014; Prioux and Barbieri, 2012); and, second, remarriage rates significantly decline with age (Smith et al., 1991; Wu et al., 2014).

In our primary widowhood duration definition, we also assume a two-year age difference between spouses ($\Delta = 2$), which corresponds to the observed median (mean is 2.5 years), for French married couples (Papon, 2019). More precisely, the numerator is the cumulated sum from 60 to ω , calculated as the product of three different terms: (1) the number of men who were two years older than women and died at the given age; (2) the number of female survivors at this age; and (3) the female life expectancy at this age. The denominator is computed as the sum from age 60 to ω of the product of (a) the number of male deaths at the considered age, and (b) the number of female survivors at this age.

$$WD^{f} = \frac{\sum_{a=60}^{\omega} d_{a+2}^{m} \cdot \frac{S_{a}^{f} + S_{a+1}^{f}}{2} \cdot \frac{e_{a}^{f} + e_{a+1}^{f}}{2}}{\sum_{a=60}^{\omega} d_{a+2}^{m} \cdot \frac{S_{a}^{f} + S_{a+1}^{f}}{2}}$$
(9)

For males, the equation is similar, with the replacement of f by m. It is important to note that, in these widowhood duration calculations for randomly paired couples, we have so far assumed no correlation in the mortality of spouses, although we will do so later in the paper.

Simulations based on Gompertz Distributions

To illustrate how changes in the shape of mortality distributions lead to shifts in the overlap between male and female distributions, subsequently affecting probability of widowhood and, ultimately, widowhood duration, we utilize Gompertz mortality distributions calibrated to INSEE's life tables. The Gompertz law is particularly well-suited to our purposes for two reasons. First, this probability law is widely used for modeling human mortality (Olshansky and Carnes, 1997). Online Appendix B.1 illustrates the quality of the fit to our data.⁵ Second, Gompertz law parameters can be easily estimated and used to compare survival curves of different populations over time. For all ages $x \ge 0$ the density of a Gompertz distribution is:

$$f_g(x,a,b) = a.exp\left[bx - \frac{a}{b}(e^{bx} - 1)\right].$$
(10)

If we set $u = \frac{1}{b} ln(\frac{b}{a})$ and $c = \frac{1}{b}$, the Gompertz density can be rewritten as a location-scale mortality density (Basellini et al., 2019):

$$f(x,c,u) = \frac{1}{c} exp\left[\frac{x-u}{c} - exp\left(\frac{x-u}{c}\right) + exp\left(-\frac{u}{c}\right)\right],\tag{11}$$

where $u \in \mathbb{R}$ is the location parameter, and c > 0 is the scale parameter of the distribution. An increase in the location parameter u shifts the mortality distribution to the right, corresponding to the well-documented increase in the modal age at death over the past decades. The scale parameter c represents the variability in the distribution. Changing

 $^{^5\}mathrm{To}$ allow the replicability of our results, we have included all parameter estimates used in this paper in Table B.1.

the scale parameter affects how far the probability distribution stretches out. A decrease in c may be considered as a compression of age at death distribution around the mode.

We further introduce an age-gap between spouses. To do so, we randomly pair 100,000 females and 100,000 males into different sex couples whose respective ages at death were drawn from Gompertz laws, and we assign them an age gap drawn from a Normal distribution. Individual widowhood duration is equal to the difference between the spouse's respective date of death plus the age gap between spouses.⁶

Finally, indrawing from research in the actuarial literature, we introduce the dependence of spousal mortality by employing a bivariate Gaussian copula (Kaishev et al., 2007). Let T_1 and T_2 be two random variables (female and male mortality distributions), and $F_1(t_1) = P(T_1 \leq t_1)$ and $F_2(t_2) = P(T_2 \leq t_2)$ be their distribution functions. The joint distribution is $F(t_1, t_2) = P(T_1 \leq t_1, T_2 \leq t_2)$. According to Sklar (1959), there must exist a copula C such that:

$$F(t_1, t_2) = C(F_1(t_1), F_2(t_2)),$$
(12)

and if F_1 and F_2 are continuous, C is unique. In practice, we use a bivariate Gaussian copula:

$$C(u_1, u_2) = P(U_1 \le u_1, U_2 \le u_2)$$
(13)

$$= P(T_1 \le \Phi^{-1}(u_1), T_2 \le \Phi^{-1}(u_2) \cup T_2 \le t_2)$$
(14)

$$= \Phi_2 \left(\Phi^{-1}(u_1), \Phi^{-1}(u_2); \rho \right), \tag{15}$$

where Φ is the cumulative distribution function (CDF) of a standard normal distribution and $\Phi_2(\rho)$ is the joint CDF of (T_1,T_2) . We utilize a Gaussian copula because it allows for a straightforward representation of the rank correlation between the two considered distributions (Klaassen and Wellner, 1997). This enables us to easily determine the sequence in which the members of the couple pass away. The Gaussian copula has a parameter ρ controlling the strength of dependence. More precisely, we first draw an age at death in a Gompertz distribution for 100,000 females and independently do the

⁶We calculate unconditional widowhood duration for the population under consideration by averaging individual widowhood duration values; any negative widowhood duration values (corresponding to cases where the wife dies before the husband) are adjusted to zero. Widowhood duration conditional on becoming a widow is computed as the average of strictly positive individual widowhood duration values.

same for 100,000 males. We then rank the two populations according to their age at death. We then use the simulated marginal distributions $F(t_1, \infty)$ and $F(\infty, t_2)$ of the Gaussian copula to determine how ranks are correlated and pair the 100,000 females and 100,000 males according to this correlation pattern. Finally, we compute widowhood duration as the difference between the matched spouse's respective dates of death and their corresponding age gap (drawn from the Normal distribution).

Results

Trends in Widowhood Duration

In France, in 2020, the expected widowhood duration at age 60 for women stood at 13 years (Figure 1) assuming a two-year age gap with their spouse. The unconditional widowhood duration was lower since it was calculated for all women, regardless of whether they outlived their husbands or not (see Background section). In 2020, the expected unconditional widowhood duration at age 60 was 9 years. The probability of widowhood stood at 70%, and is consistent with Bergeron-Boucher et al. (2022). From 1962 to 2020, widowhood duration, unconditional widowhood duration and the probability of widowhood but its duration is significant. In 2020, the expected male widowhood duration at age 60 reached 9 years. Notably, this duration has increased slightly by one year since 1962. Online Appendix B.2 shows that our results hold when we compute widowhood duration at age 50 instead of age 60 or when we switch from a period to a cohort analysis.⁷

The Role of Widowhood Duration Determinants

We now employ simulations to illustrate and quantify the role of widowhood duration determinants. Our approach unfolds progressively as we first examine the overlap of mortalities, then introduce an age gap between spouses, and, finally, incorporate dependence of spousal mortality.

We begin with a scenario featuring no age gap or dependence of mortality between

⁷The expected widowhood duration at age 50 reaches 14.7 years for females and 10.4 years for males (Appendix Figure B.2.a). Appendix Figure B.2.b displays the evolution of the expected widowhood duration at age 60 across birth cohorts (with a two-year age gap between spouses). Females born in 1902 have an expected widowhood duration of 13.6 years, while their male counterparts have a widowhood duration of 8.9 years.

spouses. Figure 2 provides a graphical assessment of the role played by overlapping mortalities in widowhood duration values. Panel (a) displays female and male mortality distributions and the degree of the overlap. Panel (b) depicts the probability of widowhood in accordance with the Equation (6) denominator, and it represents the difference between the area under the male mortality curve and the area under the male mortality curve multiplied by the female survival curve. Finally, panel (c) illustrates the unconditional widowhood duration, corresponding to the area between the joint survival and female life expectancy curves, as per the Equation (6) numerator. To evaluate the impact of mortality overlap, we performed four simulations, each corresponding to the four figures in each panel (a) to (c) of Figure 2. In these simulations, we progressively adjusted the scale and location parameters for the female distribution to match their actual values. From left to right, the first figure depicts the scenario in which women have the same mortality pattern as men (with identical location and scale parameters). In the second and third figures, we assume that only the location parameter (in the second figure) and only the scale parameter (in the third figure) of the Gompertz law for women is the same as that for men. Finally, in the fourth figure, women and men have their respective actual mortality patterns. If females and males shared the same mortality patterns, then using male mortality rates in 2020 as a reference (first figure, panel (a)) would result in perfect overlap. In this scenario, females would be equally likely to die first or become widows, resulting in a 50% probability of widowhood (first figure, panel (b)). Given the survival curve characteristics in 2020 (location parameter of 86.8 years and scale parameter of 9.6), the unconditional widowhood duration would be 5.4 years (first figure, panel (c)). Consequently, the expected widowhood duration at age 60, conditional on being a widow, would be 10.9 years. Empirically, females have a higher life expectancy that is characterized by a right-shifted mortality distribution compared to men (with a greater location parameter in the Gompertz law) and a more compressed distribution (lower scale parameter in the Gompertz law). The compression of the female distribution does not alter the probability of widowhood (second figure, panel (b)) but slightly reduces the unconditional widowhood duration to 5.2 years (second figure, panel (c)). On the other hand, shifting the female distribution to the right (third figure, panel (a)) increases both the probability of widowhood (63%, third figure, panel (b)) and the unconditional widowhood duration (8.1 years), resulting in a total widowhood duration of 12.9 years (third

figure, panel (c)). Finally, considering both the current scale and location parameters of the female mortality distribution, the probability of widowhood reaches 64% (fourth figure, panel (b)), unconditional widowhood duration reaches 7.9 years (fourth figure, panel (c)), and the total amounts to a widowhood duration of 12.4 years.⁸ In summary, it appears that only females' higher modal age at death contributes to the increase in widowhood duration compared to a scenario in which females and males share the same mortality patterns in 2020.

Figure 3 shows the sensitivity of widowhood duration both to the age gap between spouses and to the dependence of spousal mortality. In panel (a), we represent widowhood duration for different values of the age gap, drawn from a Normal distribution with mean ranging from 0 to 4 and a standard deviation of 5, aligning with the findings of Bouchet-Valat (2019). In 2020, the expected widowhood duration at age 60 for women is 13.4 years, assuming spouses have an average age gap of 2.5 years and a standard deviation of 5 for age gaps in the population.⁹ Widowhood duration shows a positive correlation with the average age gap in the population. Given the characteristics of female and male mortalities in 2020, adding one year to the age gap between spouses has a marginal effect of contributing 0.36 years to widowhood duration.¹⁰

In panel (b), we vary the degree of dependence between females and males mortalities (ρ parameter of the Gaussian copula) from 0 (independence) to 1 (perfect correlation). When $\rho=0$, widowhood duration equals 13.4 years, aligning with our previous results. If we assume a perfect correlation ($\rho=1$), widowhood duration reduces to the simple difference between average female life expectancy and average male life expectancy, as this corresponds to the scenario in which every woman dies after her husband.¹¹ The stronger the dependence of spousal mortality, the lower the widowhood duration (Figure 3b). For $\rho = 0.6$ (Kaishev et al., 2007), widowhood duration is 10.4 years.

Figure 4 summarizes the influence of these different determinants. In 2020, the ex-

⁸Note that the different methodology results in a slight difference in widowhood duration compared to the previous section (Figure 1 shows widowhood duration in 2020 to be 12.5 years, not 12.4 years). Online Appendix C provides a detailed description of the reasons for these minor discrepancies.

⁹Online Appendix B.3 provides an intermediate step in which we exclude age gap dispersion. Widowhood duration would then be only 13 years.

¹⁰Neglecting the dispersion in the age gap distribution would have led to the conclusion that the age gap plays a lesser role in widowhood duration: the marginal effect of adding one year to the age gap is equal to 0.25, consistent with life table computations.

¹¹See online Appendix B.3, panel (c) for widowhood duration without any age gap between spouses. For $\rho=1$, widowhood duration equals 4.6 years, representing the difference between female life expectancy at age 60 (27.8 years in 2020) and male life expectancy at age 60 (23.2 years in 2020).

pected widowhood duration at age 60 is 10.4 years. If we disregard the dependence of spousal mortality, widowhood duration would increase by 3 years, reaching 13.4 years. Further, if we ignore the presence of an age gap between spouses, widowhood duration would decrease by 1 year to 12.4 years. When considering that the dispersion of ages at death for females is equal to that for males, widowhood duration would increase by 0.5 years to 12.9 years. Finally, if we additionally assume that females' mortality pattern matches that of males¹², widowhood duration would be 10.9 years, representing a 2-year decrease.

Projections and Heterogeneity Results

The increase in life expectancy throughout the 20th century can be attributed to both mortality compression and the shift of mortality to older age groups (Bergeron-Boucher et al., 2015; de Beer and Janssen, 2016).¹³ These crucial demographic factors are expected to continue evolving into the future. A study conducted by Basellini and Camarda (2019), focusing on Denmark, France, Japan, and Sweden, predicts a rise in the modal age at death and a reduction in lifespan variability. These changes may, in turn, impact widowhood duration. In this section, we project widowhood duration from 2020 to 2070 using INSEE projected life tables. Figure 5 illustrates the projected evolution of widowhood duration over this time span. Female widowhood duration is forecasted to decline from 10.4 years in 2020 to 9.2 years in 2070. Subsequently, the overlap between female and male mortalities increases as a result from two key factors: the compression of mortality rates in both females (with the scale parameter expected to decrease from 8.1 to 7.6) and males (with the scale parameter expected to decrease from 9.6 to 7.7); and the shift toward older ages in both genders (with the location parameters expected to increase from 91.7 to 95.8 for females and from 86.8 to 93.7 for males).¹⁴ The probability of widowhood is projected to decrease by 7pp (-9%) from 75% to 68% between 2020 and 2070, albeit to a lesser extent than the unconditional widowhood duration, which is expected to decrease by 20% from 7.9 to 6.3 years during the same period. These findings remain

¹²The same modal and dispersion of ages at death.

¹³Mortality compression was observed during the first half of the twentieth century in Sweden Bergeron-Boucher et al. (2015) and in the Netherlands, de Beer and Janssen (2016). In the second half of the century, this dynamic was primarily replaced by the shifting of the mortality schedule, while lifespan variability remained relatively constant.

¹⁴Online appendix B.4 provides graphical illustrations for the evolution of the overlap of spousal mortalities and its role in widowhood duration.

robust across various simulation scenarios.¹⁵ Male widowhood duration is projected to experience a slight increase over this period, rising from 5.8 years to 6.2 years.

Numerous recent studies have examined the relationship between income and mortality, building upon the seminal work by Chetty et al. (2016), which uncovered a clear and large correlation between household income and individual life expectancy. Our study documents, for the first time, the correlation between household income and survivors' life expectancy while conducting heterogeneity analyses based on standards of living. Blanpain (2018) reports that in France in 2016, there is a substantial 13-year age gap in life expectancy between the first and last ventiles of the standard of living distribution for males. For females, the gradient is somewhat lower, at 8 years. We quantify here how these differences in individual life expectancy translate into widowhood duration. Figure 6 illustrates that widowhood duration decreases linearly with standard of living. Females belonging to the bottom of the standard of living distribution have an expected widowhood duration of 11.1 years at age 60, while expected widowhood duration is only 9.4 years for females at the top of the distribution. Online Appendix B.6 reveals the increase in the overlap of mortality distributions along the household income distribution, as a results of the gains in terms of modal age at death and the compression of mortality distributions. Male widowhood duration also exhibits a negative gradient along the household income distribution. Figure 6 demonstrates that men at the bottom of the income distribution have an expected widowhood duration of 5.9 years at age 60, while those at the top of the distribution have only 5.4 years.

The comparative advantage of widows from the bottom of the distribution receiving more survival benefits during their extended time as widows is significantly offset by the economic risks associated with widowhood. By adopting a perspective that allows us to examine differential mortality, Figure 6 also illustrates disparities in the joint survival of couples across the income distribution. The figure reveals that couples in the highest income group can anticipate a joint life expectancy of 23.5 years at age 60, whereas

¹⁵In addition to running simulations using the extreme life expectancy scenarios provided by INSEE, we conducted simulations in which (1°) the age gap between spouses and (2°) the dependence of spousal mortality vary by $\pm 10\%$ around the central scenario (online Appendix B.5). With the exception of the high life expectancy scenario, in which widowhood duration is forecasted to increase very slightly (up to 11 years), the anticipated small decrease in female widowhood duration remains consistent. In terms of magnitude, widowhood duration is expected to remain relatively high in 2070, ranging from 7 years (under scenarios of low life expectancy, low age gap, and high dependence) to 13.4 years (in scenarios of high life expectancy, high age gap, and low dependence).

those in the lowest income group can only expect 16.8 years of shared life together. The increased risk of poverty among widows may also stem from greater variability in widowhood duration within the lowest income segment of the population. An extended widowhood duration amplifies the risk of poverty due to two key factors: (i) the higher likelihood of adverse events occurring over an extended time frame, and (ii) the additional impact of retirees experiencing a decline in their standard of living over time compared to those still in the labor market.¹⁶ The average widowhood duration alone does not provide insights into its dispersion. In Figure 7 we present the proportion of widows whose widowhood duration is expected to exceed 15, 20, or 25 years per standard of living ventile. Among widows in the lowest income ventile, 28% are expected to have a widowhood duration exceeding 15 years, whereas this is the case for only 19% in the highest ventile. Furthermore, 4.5% of widows in the lowest standard of living ventile are likely to be widows for more than 25 years, compared to only 2.1% in the highest ventile.

Discussion

In this paper, we estimate the factors influencing widowhood duration within a unified framework. We show that, in France in 2020, the expected widowhood duration for females at age 60 is 10.4 years, which depends on three primary determinants: *(i)* the mortality distributions of both females and males, including their degree of overlap; *(ii)* the age gap between spouses; and *(iii)* the dependence of spousal mortality. Starting from our most precise definition of widowhood duration, we sequentially isolate each determinant down to the most basic definition, thus quantifying their role in the measurement. First, consider the dependence of spousal mortality, a factor not previously accounted for in the literature on either widowhood duration (Compton and Pollak, 2021) or the probability of widowhood (Bergeron-Boucher et al., 2022). This dependence reduces widowhood duration by 3 years. Had we assumed that mortality distributions of spouses are independent, widowhood duration would have been 13.4 years. Furthermore, we extend the work of Goldman and Lord (1983), which focused on specific age gap combinations. Our extension reveals that, when considering the entire age gap distribution among spouses in the population, widowhood duration increases by 1 year, with 0.4 years

 $^{^{16}\}mathrm{Pension}$ benefits do not increase as fast as the average income in the economy because they are indexed to inflation.

attributed to dispersion around the mean. If we had assumed the same age for spouses, widowhood duration would have been 12.4 years. And if the dispersion of ages at death for females is the same as for men, widowhood duration would increase by 0.5 years, yielding 12.9 years. Finally, assuming that the female mortality pattern matches that of males, not only having the same dispersion of ages at death but also the same modal age at death, would result in a widowhood duration of 10.9 years (a reduction of 2 years).

Our methodology uses simulations based on the Gompertz law and a bivariate Gaussian copula. Specifically, we use the INSEE life tables as our primary data source (Compton and Pollak, 2021). Building upon the work of Vaupel et al. (2021), we introduce a temporal dimension to their outsurvival indicator. We calculate the probability of widowhood, unconditional widowhood duration, and widowhood duration conditional on being widowed using a Gompertz model transformed into a location-scale mortality density (Basellini et al., 2019). With this approach, we can directly analyze the role of compression and right-shifting of mortality in the widowhood duration measure. The flexibility of our simulation approach allows us to generate a population of couples with age gaps drawn from observed population distributions (Bouchet-Valat, 2019) and to correlate death ranks among spouses using a Gaussian copula.

Using this method, we derive two sets of novel results that shed light on three important policy questions in the French context and potentially in other contexts when individuals or nations plan for the long-term needs of aging populations. First, using INSEE projected life tables for the period 2020–2070, we demonstrate that widowhood duration is expected to slightly decrease over this period but will remain high. By 2070, the expected widowhood duration is projected to be 9.2 years for females at age 60 and 6.3 years for men. This underscores the importance for policymakers to recognize that even with the convergence of life expectancies between females and males, widows and widowers may still experience periods of vulnerability. Given the expected increase in the number of widows as the baby boomers reach the age of widowhood, policymakers in our aging societies will need to address this large-scale issue in the coming years. In particular, our unconditional widowhood duration indicator offers a valuable measure of the overall level of widowhood protection required by a population. It estimates the average duration during which a widow or widower may be eligible for social benefits, considering that some individuals may not experience widowhood. When applied to the population of couples aged 60 and over, this indicator should help policymakers to determine how many people, for whom to provide support and for how long, in the event of widowhood. Such support may include public survivors' benefits and formal long-term care facilities, as the deceased spouse has often been the primary caregiver. Second, our analysis of widowhood duration heterogeneity based on standard of living reveals that females at the bottom of the distribution experience a longer expected widowhood duration compared with females at the top. This result is of vital interest for public policy, as individuals tend to underestimate the likelihood of outliving their spouse. These expectations may vary by income level, with the literature indicating that individuals with low socioeconomic status have a lower subjective probability of survival compared with those with high socioeconomic status (Delavande and Rohwedder, 2011). This misperception can lead to inadequate savings – reinforced by lower financial literacy among the less educated (Lusardi and Mitchell, 2023) – potentially resulting in poverty for widows, particularly those with already limited resources. Furthermore, the dispersion of widowhood duration is greater among individuals with lower standards of living, which amplifies their risk of poverty. Additionally, our findings indicate a dual penalty associated with differential mortality. Low-income females not only have shorter lifespans but also spend a larger proportion of their life expectancy in widowhood, leading to increased experiences of loneliness. Third, by comparing widowhood duration with the amount of benefits paid across the income distribution, it is possible to draw up precise budgets for survivors' benefits. This is particularly useful for designing survivor benefit programs, such as introducing a minimum pension or implementing a means test. According to our results, introducing a minimum pension would reduce the long-term poverty risk for individuals in long-term low-resource situations. To maintain a constant budget for survivor benefits, a means test would need to be applied to those with the highest incomes. The goal is to avoid compensating individuals with substantial resources, and maybe also private insurance, at the expense of those with less coverage or greater vulnerability. Furthermore, our approach can be applied to other settings with two correlated survival times. It could shed light on other policy-relevant scenarios, such as long-term care policy, by providing information on the duration of informal spousal long-term care.

At this stage, we do not address three limitations discussed in the literature. First, we use life tables computed for the entire population, without using life tables specific to marital status. Consequently, we cannot account for the fact that mortality among married individuals tends to be lower than that of singles, possibly due to healthier individuals self-selecting into marriage (Goldman, 1993) and to the protective effects of marriage (Sanders and Melenberg, 2016). Second, since our focus is on changes in widowhood duration over time, most of our measures are based on period mortality rates. While we observed similar patterns in average widowhood duration across birth cohorts, stability in age-specific mortality rates, upon which most of our results rely, is highly unlikely. Third, the life tables we use do not differentiate between migrants and natives, so we cannot assess how widowhood duration varies between these groups, whose mortality patterns differ (Guillot et al., 2023). Furthermore, we are unable to quantify the extent to which the increasing migrant population in France might contribute to changes in widowhood duration over time. Investigating this issue in countries where migrant-specific data are accessible would be particularly useful. The method employed in this paper also has three inherent limitations. First, we do not present a comprehensive global decomposition of widowhood duration. Instead, we examine the impact of each determinant individually while holding all other parameters constant at their central values. In doing so, we do not account for potential endogeneity among these determinants, such as the documented influence of age gap on mortality (Drefahl, 2010). Second, the likelihood of both the age gap between spouses and the level of spousal mortality dependence remaining constant is improbable across the ventiles of standard of living (Dabergott, 2022). If we had assumed a larger age gap at the bottom of the income distribution and a higher mortality correlation at the top of the income distribution, our measure of the widowhood income gradient might have been slightly higher. However, we currently lack the empirical evidence needed to refine our hypotheses. This issue can be addressed using administrative data, which we reserve for future research. In any case, the bias does not undermine our findings, as our results should be seen as a lower bound on the income gradient in widowhood duration. Third, our calculations are based on two assumptions regarding partnership behaviors that align with the current predominant behaviors among individuals aged 60 and over. Divorce after age 60 remains rare (Solaz, 2021), and repartnering after widowhood is still uncommon (Bonnet et al., 2019). However, these behaviors may become more prevalent in the future, as the incidence of divorce among those aged 60 and older is increasing (the so-called grey divorce revolution; see Brown and Lin, 2012), and

repartnering after widowhood may become more common at older ages. Our approach may overestimate the survivor benefits widows would claim in the future and the duration for which they would receive them.¹⁷ However, from an individual point of view, these calculations are valuable for assessing one's vulnerability in widowhood. Less well-off widows are not more likely to repartner than others (Bonnet et al., 2019), confirming concerns about longer widowhood durations and the associated poverty risks for widows at the bottom of the income distribution.

Under the same hypotheses as Compton and Pollak (2021), our results suggest that widowhood duration in France is slightly lower in 2020 than in the US, though still within the same range. This difference arises because widowhood duration in France has stabilized since the 1990s, while it exhibits a U-shaped pattern over time in the US. This last observation highlights the need for a more systematic international comparison, a potential avenue for future research.

¹⁷In France, survivor benefits may be reduced or eliminated in cases of divorce and can be suspended if the widow remarries after widowhood.

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Figures



Figure 1: Evolution of widowhood duration over the years Note: We assume a fixed 2-year age gap between spouses. Source: INSEE life tables.

0.06 0.06 0.06 0.06 Density 0.04 Density 0.04 0.04 0.04 Density Density 0.02 0.02 0.00 0.00 0.00 0.00 80 120 60 80 100 120 80 100 120 60 100 60 100 60 80 120 Age Age Age Age ---- M ---- M F = loc M ····· M F = scale M ---- I F = M (b) Widowhood probability 0.04 0.04 0.04 0.04 0.03 0.03 0.03 0.03 Probability Probability Probability Probability 0.02 0.02 0.02 0.02 0.01 0.01 0.01 0.0 0.00 0.00 0.00 0.00 60 100 80 100 80 100 80 100 120 80 120 120 60 60 120 60 Age Age Age Age dm --- Sf x dm dm Sf x dm dm 🔤 Sf x dm dm ----Sf x dm (c) Unconditional widowhood duration 1.00 1.00 1.00 1.00 0.75 0.75 0.75 0.75 QM0.50 QM 0.50 QM0.50 UWD 0.50 0.25 0.25 0.25 0.25 0.00 0.00 0.00 0.00 100 100 100 120 60 80 120 120 100 80 60 80 60 80 120 60 Age Age Age Age Sf ---- Sm x Sf Sf --- Sm x Sf Sf ---- Sm x Sf Sf ---- Sm x Sf

(a) Mortality distribution

Figure 2: The role of mortality overlap in widowhood duration, in 2020 Note: Mortality distributions are simulated using a Gompertz law and assuming neither an age gap between spouses nor any dependence of spousal mortality. UWD stands for unconditional widowhood duration. M=F stands for the scenario in which women have the same mortality pattern as men. F=loc M stands for the scenario in which we assume that only the location parameter of the Gompertz law for women is the same as that for men. F=scale M stands for the scenario in which we assume that only the scale parameter of the Gompertz law for women is the same as that for men. Source: Authors' simulations and INSEE life tables.



Figure 3: Widowhood duration in relation to age gap and dependence of spousal mortality Note: Mortality distributions are simulated using a Gompertz law. Source: Authors' simulations and INSEE life tables.



5. Widowhood duration without (a,b), with (c), and with (d) men's modal age at death

Figure 4: Determinants of female widowhood duration in 2020

Note: In 2020, expected widowhood duration at age 60 is 10.4 years. Ignoring the dependence of spousal mortality would have increased widowhood duration by 3 years (Δ), to 13.4. Disregarding the existence of an age gap between spouses would have reduced it by 1 year, to 12.4. Considering equal dispersion of ages at death between females and males would have increased widowhood duration by 0.5 years, to 12.9. Finally, assuming that the female mortality pattern matches that of males (same modal and dispersion of ages at death) would result in a widowhood duration of 10.9 years (2 years lower). Mortality distributions are simulated with a Gompertz law.

Source: Authors' simulations and INSEE life tables.



Figure 5: Projected evolution of widowhood duration over the years Note: Mortality distributions are simulated using a Gompertz law. Source: Authors' simulations and INSEE life tables.



Figure 6: Widowhood and joint survival durations per standard of living in 2016 Note: Mortality distributions are simulated using a Gompertz law. Source: Authors' simulations and INSEE life tables.



Figure 7: Median and extreme widowhood duration by standard of living (females) Note: Mortality distributions are simulated using a Gompertz law. Source: Authors' simulations and INSEE life tables.