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Constructing prediction intervals for the age distribution of deaths

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ABSTRACT

We introduce a model-agnostic procedure to construct prediction intervals for the age distribution of deaths. The age distribution of deaths is an example of constrained data, which are nonnegative and have a constrained integral. A centered log-ratio transformation and a cumulative distribution function transformation are used to remove the two constraints, where the latter transformation can also handle the presence of zero counts. Our general procedure divides data samples into training, validation, and testing sets. Within the validation set, we can select an optimal tuning parameter by calibrating the empirical coverage probabilities to be close to their nominal ones. With the selected optimal tuning parameter, we then construct the pointwise prediction intervals using the same models for the holdout data in the testing set. Using Japanese age- and sex-specific life-table death counts, we assess and evaluate the interval forecast accuracy with a suite of functional time-series models.

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Compositional data analysis; functional principal component analysis; functional time series; prediction interval calibration; split conformal prediction; standard deviation-based conformity

1. Introduction

Actuaries and demographers have long been interested in developing statistical techniques to model and forecast mortality for annuity pricing and government planning. In the literature on human mortality, three functions are widely studied: mortality rate, survival function, and life-table death counts (representing age distribution of deaths). Although these three functions are complementary (see, e.g. Dickson et al., 2009; Preston et al., 2001), they differ by the number of constraints. The mortality rate is between 0 and 1; the survival function is also between 0 and 1 and exhibits monotonicity over a certain age group; and the life-table death counts are non-negative and sum up to a radix, commonly 10^5 .

Most of the literature has focused on the development of novel approaches for modeling and forecasting age-specific logarithmic mortality rates (see, e.g. Basellini, Camarda et al., 2023; Booth, 2006; Booth & Tickle, 2008, for comprehensive reviews). Instead of modeling central mortality rates, we consider modeling life-table deaths as an example of a probability density function (see, e.g. Basellini, Kjaergaard et al., 2020). Observed over a period, we could visualize, model, and forecast a redistribution of the life-table deaths, where deaths at younger ages are shifted gradually toward older ages due to longevity. In addition to providing an informative description of the mortality experience of a population, life-table deaths provide readily available information on central longevity indicators

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(see, e.g. Canudas-Romo, 2010; Cheung et al., 2005) and lifespan variability (see, e.g. Aburto & van Raalte, 2018; Aburto et al., 2020; Robine, 2001; van Raalte & Caswell, 2013; Vaupel et al., 2011).

To model the age distribution of deaths, we resort to an extrinsic approach via transformation. In demography, Bergeron-Boucher, Canudas-Romo et al. (2017) and Bergeron-Boucher, Simonacci et al. (2018) apply the centered log ratio (CLR) transformation to obtain unconstrained data, which can then be modeled through principal component analysis. In actuarial science, Shang and Haberman (2020) and Shang et al. (2022) used the forecasted life-table death counts to calculate estimated fixed-term annuity prices. In statistics, Stefanucci and Mazzucco (2022) apply the CLR transformation to model cause-specific mortality data, Delicado (2011) apply the CLR transformation to analyze density functions over space, and Kokoszka et al. (2019) model and forecast financial time series of density functions.

An issue with the CLR transformation is the presence of zero counts. Some ad-hoc ways of handling zero counts exist, including adding or subtracting a small constant (see, e.g. Fry et al., 2000; Martín-Fernández et al., 2013). Recently, Shang and Haberman (2025) introduced a cumulative distribution function (CDF) transformation with the advantage of monotonicity. We first normalize the age distribution of death so that the radix is one, akin to the probability density function (PDF), and then convert the PDF to a CDF. The inverse of CDF is quantile, which is a key quantity in the Wasserstein distance to measure the discrepancy between two distributions (see, e.g. Dubey & Müller, 2022). With a time series of CDFs, we model its pattern via a logistic transformation. Within this unconstrained space, we apply a suite of functional time-series forecasting methods to obtain the h -step-ahead curve prediction for a chosen forecast horizon h . By taking the inverse logistic transformation, the h -step-ahead forecast life-table death counts are obtained after first-order differencing and renormalized to the original scale.

The current literature lacks guidance on the construction of prediction intervals for the age distribution of deaths. We aim to present a general procedure that works for time-series forecasting models in Section 3. The general procedure divides the data samples into training, validation, and testing sets. The validation set allows us to tune an optimal parameter that adjusts the prediction intervals so that the empirical coverage probability is close to its nominal one. With the selected optimal parameter, we construct the pointwise prediction intervals for the data in the holdout set. Assuming that the data in the validation and testing sets do not differ much, our construction can achieve satisfactory coverage. Using age- and sex-specific life-table death counts in Japan in Section 2, we study the interval forecast accuracy of several functional time-series methods in Section 4. The conclusion is presented in Section 5, along with some ideas on how the methodology can be further extended.

2. Period life-table death counts

In many developed countries, such as Japan, increases in longevity risk and an aging population have led to concerns about the sustainability of government pension, health and age care systems. Japan has one of the highest average life expectancies in the world, with extreme longevity in Okinawa prefecture (Coulmas, 2007).

Our chosen mortality instrument is the life-table death counts, where the life-table radix is fixed at 100,000 at age zero while the remaining number of people alive is 0 in the last age group 110+ for each year. There are 111 ages, which are 0, 1, ..., 109, 110+. Due to rounding, there are potentially zero counts for people aged 110+ at some years. To overcome this problem, we work with the probability of dying (i.e. q_x) and the radix of the life table to recalculate our estimated death counts (up to six decimal places). In doing so, we obtain more precise death counts than those reported in Japanese Mortality Database (2025).

In Figure 1, we present Japanese age- and sex-specific life-table death counts from 1975 to 2022, obtained from Japanese Mortality Database (2025). We used data from the period after the First and Second World Wars to obtain a more stable parameter estimate from the historical

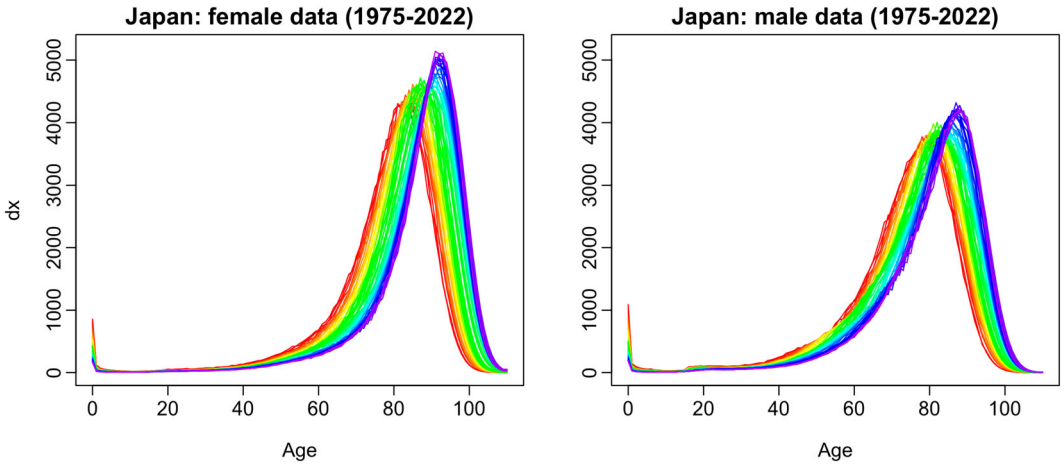


Figure 1. Rainbow plots of the age distribution of deaths from 1975 to 2022 in a single-year group. The life-table radix is 100,000 for each year. The life-table death counts in the oldest years are shown in red, while the most recent years are in violet. Curves are ordered chronologically by the rainbow colors.

data. The beginning year 1975 was chosen to be the same as its subnational data (see also Shang & Haberman, 2025).

Figure 1 demonstrates a decreasing trend in infant death counts and a typical negatively skewed distribution for life-table death counts, where the peak shifts to higher ages for both sexes. This shift is due to the risk of longevity, which concerns insurers and pension funds in transferring and managing the risks of annuity products (see Denuit et al., 2007, for a discussion). By modeling the period life-table death counts, we can understand a redistribution of life-table death counts, where deaths at younger ages gradually shift towards older ages.

Since the exposure-to-risk can be difficult to estimate accurately due to migration, under-reporting, or late registration (see, e.g. Cairns et al., 2016), we choose to work with life-table death counts instead of mortality rates. Life-table death counts are derived from the probability of dying and bypass the need for direct exposure estimation, and they represent the number of deaths in the implied stationary population and lead to the corresponding probability density function of the age distribution of deaths. Due to non-negativity and summability constraints, we can study the age distribution of deaths for all available ages.

3. Construction of prediction intervals for density-valued objects

Denote age-specific life-table death counts as $d_t^s(u)$, where t denotes a year, s denotes female or male data, and u represents an age. For each year t , the life-table death counts sum to a radix 10^5 . In Sections 3.1 and 3.2, we consider two transformation methods to remove constraints in the life-table death counts. For modeling the unconstrained data within each transformation, we consider three functional time-series forecasting models in Section 3.3.

3.1. Centered log-ratio transformation

By treating age as a continuous variable, the CLR transformation can be written as

$$\text{CLR}[d_t^s(u)] := G_t^s(u) = \ln d_t^s(u) - \frac{1}{\eta} \int_u \ln d_t^s(u) du,$$

where η denotes the length of the age interval and $\frac{1}{\eta} \int_u \ln d_t^s(u) du$ is the geometric mean. With a time series of functions $[G_1^s(u), \dots, G_n^s(u)]$, we apply the univariate, multivariate, and multilevel functional

time-series models to obtain h -step-ahead forecasts $\widehat{G}_{n+h|n}^s(u)$. A brief description of these time-series models is given in Section 3.3. Via the inverse CLR transformation, the forecast life-table death counts can be expressed as

$$\widehat{d}_{n+h|n}^s(u) = \frac{\exp^{\widehat{G}_{n+h|n}^s(u)}}{\int_u \exp^{\widehat{G}_{n+h|n}^s(u)} du} \times 10^5.$$

3.2. Cumulative distribution function transformation

By normalizing the life-table radix from 10^5 to one, the first transformation computes the empirical CDF via cumulative sum,

$$D_t^s(x) = \sum_{u=1}^x d_t^s(u), \quad x = 1, \dots, 111.$$

Since $D_t^s(111) = 1$, we apply the logistic transformation to the first 110 ages,

$$L_t^s(y) = \ln \left[\frac{D_t^s(y)}{1 - D_t^s(y)} \right], \quad y = 1, \dots, 110,$$

where $\ln(\cdot)$ denotes the natural logarithm. With a time series of functions $[L_1^s(y), \dots, L_n^s(y)]$, we obtain h -step-ahead forecasts, denoted by $\widehat{L}_{n+h|n}^s(y)$, via the univariate, multivariate, and multilevel functional time-series methods.

Taking the inverse logit transformation, we obtain

$$\widehat{D}_{n+h|n}^s(y) = \frac{\exp^{\widehat{L}_{n+h|n}^s(y)}}{1 + \exp^{\widehat{L}_{n+h|n}^s(y)}}.$$

By including a column of ones $\mathbf{1}$, we obtain $\widehat{D}_{n+h|n}^s(x) = [\widehat{D}_{n+h|n}^s(y), \mathbf{1}]$. By taking the first-order differencing, we obtain

$$\begin{aligned} \widehat{d}_{n+h|n}^s(z) &= \Delta \widehat{D}_{n+h|n}^s(z) \\ &= \widehat{D}_{n+h|n}^s(z) - \widehat{D}_{n+h|n}^s(z-1), \quad z = 2, \dots, 111, \end{aligned}$$

where Δ represents the first-order differencing, and $\widehat{d}_{n+h|n}^s(1) = \widehat{D}_{n+h|n}^s(1)$. Given the life-table radix of 10^5 , we renormalize the forecasts to their original scale: $\widehat{d}_{n+h|n}^s(u) = \widehat{d}_{n+h|n}^s(z) \times 10^5$.

3.3. A suite of functional time-series forecasting methods

The unconstrained data are assumed to be elements of the Hilbert space equipped with the inner product. We model the unconstrained data, $G_t^s(u)$ in the CLR transformation or $L_t^s(u)$ in the CDF transformation. For illustration, we demonstrate our idea with $G_t^s(u)$, which can be expressed via the Karhunen-Loève expansion as

$$G_t^s(u) = \sum_{k=1}^{K_s} \eta_{t,k}^s \psi_k^s(u) + \epsilon_t^s(u), \quad (\text{UFTS})$$

where $\psi_k^s(u)$ denotes the k^{th} functional principal component for age u and sex s , $\eta_{t,k}^s = \langle G_t^s(u), \psi_k^s(u) \rangle$ is the estimated principal component score at time t and $\langle \cdot, \cdot \rangle$ denotes the L^2 inner product, $\epsilon_t^s(u)$

denotes the residual function for age u and sex s in year t , and K_s denotes the number of functional principal components. We consider an eigenvalue ratio (EVR) criterion of Li et al. (2020) to select the number of K_s , which is the integer that minimizes the ratio of two adjacent empirical eigenvalues given by

$$K_s = \underset{1 \leq \kappa \leq (n-1)}{\operatorname{argmin}} \left\{ \frac{\widehat{\lambda}_{\kappa+1}^s}{\widehat{\lambda}_{\kappa}^s} \times \mathbb{1} \left(\frac{\widehat{\lambda}_{\kappa+1}^s}{\widehat{\lambda}_{\kappa}^s} \geq \delta \right) + \mathbb{1} \left(\frac{\widehat{\lambda}_{\kappa+1}^s}{\widehat{\lambda}_{\kappa}^s} < \delta \right) \right\}, \quad (1)$$

where $\widehat{\lambda}_{\kappa}^s$ is the κ^{th} estimated eigenvalue, δ is a prespecific small positive number, set as $\delta = 1 / \ln(\max\{\widehat{\lambda}_1^s, n\})$, and $\mathbb{1}(\cdot)$ denotes the binary indicator function. For comparison, we also consider $K_s = 6$ used in Hyndman et al. (2013).

We also consider a multivariate functional time-series method to jointly model and forecast the female and male series that could be correlated. Let $G_t^F(u)$ and $G_t^M(u)$ represent unconstrained female and male data. By stacking both series in a vector, we compute their joint covariance function. Via Karhunen-Loève expansion, a realization of both series can be approximated by

$$\mathbf{G}_t(u) = \boldsymbol{\theta}(u) + \boldsymbol{\Phi}(u)\boldsymbol{\beta}_t^{\top}, \quad (\text{MFTS})$$

where $\mathbf{G}_t(u) = [G_t^F(u), G_t^M(u)]^{\top}$; $\boldsymbol{\theta}(u) = [\theta^F(u), \theta^M(u)]^{\top}$ denotes the mean functions for the female and male series, respectively; $\boldsymbol{\Phi}(u)$ is a $(2 \times (K \times 2))$ matrix, where the off-diagonal elements capture the correlation between the estimated principal components; $\boldsymbol{\beta}_t = [\boldsymbol{\beta}_t^F, \boldsymbol{\beta}_t^M]$ and $\boldsymbol{\beta}_t^F = [\beta_{t,1}^F, \dots, \beta_{t,K}^F]$ denotes the estimated principal component scores.

The multilevel functional time-series method extracts a common pattern shared by female and male series $R_t(u)$ and a series-specific residual pattern $U_t^s(u)$. Via functional principal component analysis, the common and residual patterns are modeled by projecting the data onto the eigenfunctions of the covariance functions of aggregated and series-specific curves, respectively. For $t = 1, 2, \dots, n$, a realization can be approximated by

$$G_t^s(u) = \mu^s(u) + R_t(u) + U_t^s(u). \quad (\text{MLFTS})$$

With a finite sample, we estimate

$$\begin{aligned} \widehat{\mu}^s(u) &= \frac{1}{n} \sum_{t=1}^n G_t^s(u) \\ R_t(u) &\approx \sum_{k=1}^K \beta_{t,k} \phi_k(u) \\ U_t^s(u) &\approx \sum_{\ell=1}^V \gamma_{t,\ell} \psi_{\ell}(u), \end{aligned}$$

where K and V represent the number of functional principal components retained. These components can be determined by the EVR criterion in (1) or set to $K = V = 6$.

3.4. Construction of prediction intervals

We equally divide the data sample consisting of 48 years from 1975 to 2022 into training, validation, and testing sets, each consisting of 16 years. Using the data in the training sample, we implement an expanding window forecast scheme to obtain the h -step-ahead density forecasts in the validation set for $h = 1, 2, \dots, 15$. The expanding window scheme allows one to assess how a forecasting method performs on short and medium horizons. We have different numbers of curves in the validation set for each forecast horizon. For example, when $h = 1$, we have 16 years to evaluate the forecast errors; when

$h = 15$, we have two years to evaluate the residual functions between the samples in the validation set and their forecasts, and compute their functional standard deviation. Note that we need at least two years of data to compute the functional standard deviation. Forecast errors are denoted by $\widehat{\varepsilon}_m(u) = d_m^s(u) - \widehat{d}_m^s(u)$, for $m = 1, 2, \dots, M$, and M denotes the number of years of residual functions.

Let us define $\gamma(u) = \text{sd}[\widehat{\varepsilon}_m(u)]$. For a level of significance α , our aim is to determine $(\underline{\zeta}_\alpha, \bar{\zeta}_\alpha)$ such that $\alpha \times 100\%$ of the residuals satisfy

$$-\underline{\zeta}_\alpha \gamma(u) \leq \widehat{\varepsilon}_m(u) \leq \bar{\zeta}_\alpha \gamma(u).$$

$(\bar{\zeta}_\alpha, \underline{\zeta}_\alpha)$ are the tuning parameters mentioned in the abstract and Section 1. Typically, the constants $\bar{\zeta}_\alpha$ and $\underline{\zeta}_\alpha$ are chosen equal. By the law of large numbers, one may achieve

$$\Pr[-\underline{\zeta}_\alpha \gamma(u) \leq d_{n+h}^s(u) - \widehat{d}_{n+h|n}^s(u) \leq \bar{\zeta}_\alpha \gamma(u)] \approx \frac{1}{M} \sum_{m=1}^M \mathbb{1}[-\underline{\zeta}_\alpha \gamma(u) \leq \widehat{\varepsilon}_m(u) \leq \bar{\zeta}_\alpha \gamma(u)].$$

To determine the optimal ζ_α , the samples in the validation set are used to calibrate the prediction intervals so that the empirical coverage probabilities are close to their nominal coverage probabilities. As an output of this calibration, we obtain an optimal tuning parameter based on the coverage probability difference in Section 4.2.

For comparison, we also consider conformal prediction intervals, which are well calibrated in a large sample size (Dhillon et al., 2024). The conformal prediction introduced by Shafer and Vovk (2008) is a popular methodology in machine learning and is used to construct probabilistic forecasts calibrated on out-of-sample errors. Since its introduction in Gammerman et al. (1998), it has received increasing attention in various fields, including time series forecasting (Angelopoulos et al., 2023; Fontana et al., 2023; Yu & Xie, 2021) and climate modeling (Cannon, 2018; Qian & Chang, 2021). The conformal prediction is model-agnostic and presents a distribution-free way to construct prediction sets with a finite-sample coverage guarantee. From the absolute value of $\widehat{\varepsilon}_m(u)$, we calculate its $100(1 - \alpha)\%$ quantile for a level of significance α , denoted by $q_\alpha(u)$. The prediction interval can be obtained as

$$\left[\widehat{d}_{n+h|n}^s(u) - q_\alpha(u), \widehat{d}_{n+h|n}^s(u) + q_\alpha(u) \right],$$

where $\widehat{d}_{n+h|n}^s(u)$ denotes the h -step-ahead point forecasts for the data in the test set.

We consider the simplest conformity score by taking the quantiles from the absolute residuals. We acknowledge that other conformity scores, such as the use of quantile regression, are possible to construct asymmetric prediction intervals that may lead to better performance (see, e.g. Chernozhukov et al., 2021; Romano et al., 2019). Two limitations are commonly associated with the split conformal prediction: First, it works well for identically distributed data under the assumption of exchangeability. For time series data, the empirical coverage deteriorates as the forecast horizon increases. Second, it requires a large sample size for the validation and testing sets to achieve superior calibration.

4. Evaluation of interval forecast accuracy

4.1. Expanding-window forecast scheme

An expanding window analysis of a time-series model is commonly used to assess model and parameter stability over time. With the samples in the test set, we evaluate and assess the accuracy of the interval forecast. Using the first 32 years from 1975 to 2006, we can produce one- to 16-step-ahead forecasts. Through an expanding window scheme, we estimate the parameters in the time-series forecasting models using the first 33 observations from 1975 to 2007. Forecasts from the estimated model are produced for one- to 15-step-ahead forecasts. We iterate this process by increasing the sample size

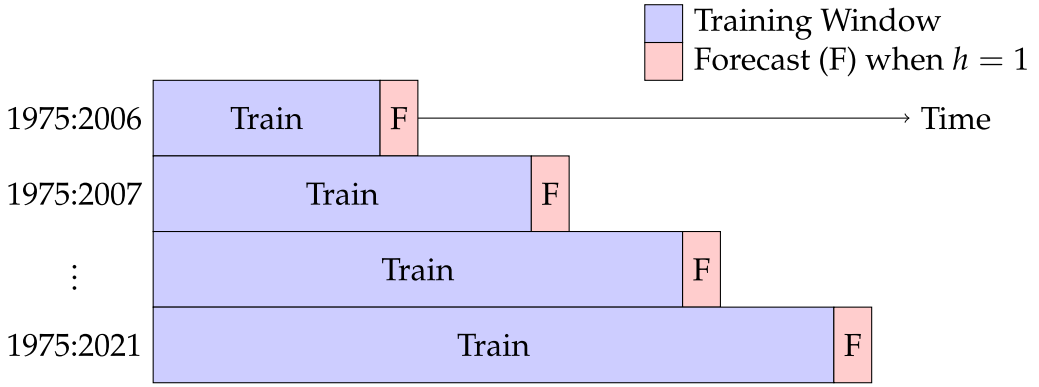


Figure 2. A diagram of the expanding-window forecast scheme.

by one year until we reach the end of the data period in 2022. This iteration process produces 16 one-step-ahead forecasts, 15 two-step-ahead forecasts, $\dots$, and one 16-step-ahead forecast. In Figure 2, we show a diagram of the expanding window forecast scheme for the forecast horizon $h = 1$, although we also consider other forecast horizons from $h = 2$ to 15.

4.2. Interval forecast errors

To evaluate interval forecast accuracy, we consider the coverage probability difference (CPD) between the empirical coverage probability (ECP) and nominal coverage probability, as well as the mean interval score of Gneiting and Raftery (2007). For each year in the forecast period, the h -step-ahead prediction intervals are calculated at the $100(1 - \alpha)\%$ nominal coverage probability. We consider the common case of the symmetric $100(1 - \alpha)\%$ prediction intervals, with lower and upper bounds that are quantiles at $\alpha/2$ and $1 - \alpha/2$, denoted by $\hat{d}_{v+\zeta}^{s,lb}(u)$ and $\hat{d}_{v+\zeta}^{s,ub}(u)$. The ECP and CPD are defined as

$$\text{ECP}_h = \frac{1}{111 \times (16 - h)} \times \sum_{\zeta=h}^{16} \sum_{u=1}^{111} \mathbb{1} \left\{ \hat{d}_{v+\zeta}^{s,lb}(u) \leq d_{v+\zeta}(u) \leq \hat{d}_{v+\zeta}^{s,ub}(u) \right\},$$

$$\text{CPD}_h = \frac{1}{111 \times (16 - h)} \times \sum_{\zeta=h}^{16} \sum_{u=1}^{111} \left[\mathbb{1} \{ d_{v+\zeta}(u) > \hat{d}_{v+\zeta}^{s,ub}(u) \} + \mathbb{1} \{ d_{v+\zeta}(u) < \hat{d}_{v+\zeta}^{s,lb}(u) \} \right],$$

where v denotes the years in the training and validation sets.

For different ages and years in the test set, the mean and median ECP are defined as

$$\overline{\text{ECP}} = \frac{1}{15} \text{ECP}_h,$$

$$\text{M}[\text{ECP}] = \text{median}(\text{ECP}_h).$$

Similarly, mean and median CPD are defined as

$$\overline{\text{CPD}} = \frac{1}{15} \sum_{h=1}^{15} \text{CPD}_h,$$

$$\text{M}[\text{CPD}] = \text{median}(\text{CPD}_h).$$

As defined by Gneiting and Raftery (2007), a scoring rule for the prediction intervals at age u is

Table 1. For different forecast horizons $h = 1, 2, \dots, 15$, we present the estimated tuning parameter ξ_α values obtained from the univariate functional time-series model (the number of retained principal components, K , can be determined via the EVR criterion or set as six) with the CDF transformation.

h	$\alpha = 0.2$				$\alpha = 0.05$			
	EVR		$K = 6$		EVR		$K = 6$	
	F	M	F	M	F	M	F	M
1	1.41	1.55	1.37	1.40	2.09	2.28	1.98	2.28
2	1.42	1.52	1.39	1.43	2.09	2.28	1.98	2.28
3	1.47	1.57	1.44	1.47	2.16	2.32	2.04	2.32
4	1.57	1.66	1.53	1.59	2.35	2.45	2.25	2.45
5	2.00	1.95	1.95	1.76	2.84	2.85	2.78	2.85
6	2.28	3.26	2.22	2.99	3.40	4.50	3.35	4.50
7	2.40	3.42	2.38	3.08	3.47	4.63	3.52	4.63
8	2.55	3.69	2.51	3.35	3.55	5.17	3.49	5.17
9	2.63	4.36	2.59	3.99	3.46	6.00	3.51	6.00
10	3.87	6.04	3.83	5.58	5.23	7.71	5.34	7.71
11	4.38	6.19	4.21	5.84	6.13	9.03	6.01	9.03
12	5.46	7.40	5.25	7.40	8.27	11.78	8.03	11.78
13	4.34	7.24	4.35	6.65	6.15	16.56	6.30	16.56
14	4.68	8.20	5.26	8.24	8.63	15.28	9.11	15.28
15	3.35	11.84	3.51	12.36	11.83	18.55	15.28	18.55

$$\begin{aligned}
S_{\alpha, \zeta} \left[\widehat{d}_{v+\zeta}^{s, \text{lb}}(u), \widehat{d}_{v+\zeta}^{s, \text{ub}}(u), d_{v+\zeta}^s(u) \right] &= \left[\widehat{d}_{v+\zeta}^{s, \text{ub}}(u) - \widehat{d}_{v+\zeta}^{s, \text{lb}}(u) \right] \\
&+ \frac{2}{\alpha} \left[\widehat{d}_{v+\zeta}^{s, \text{lb}}(u) - d_{v+\zeta}^s(u) \right] \mathbb{1} \left\{ d_{v+\zeta}^s(u) < \widehat{d}_{v+\zeta}^{s, \text{lb}}(u) \right\} \\
&+ \frac{2}{\alpha} \left[d_{v+\zeta}^s(u) - \widehat{d}_{v+\zeta}^{s, \text{ub}}(u) \right] \mathbb{1} \left\{ d_{v+\zeta}^s(u) > \widehat{d}_{v+\zeta}^{s, \text{ub}}(u) \right\},
\end{aligned}$$

where the level of significance is customarily set to $\alpha = 0.2$ or 0.05 . The interval score rewards a narrow prediction interval width if and only if $100(1 - \alpha)\%$ of the holdout densities lies within the prediction interval.

For different ages and years in the test set, the mean interval score is defined by

$$\bar{S}_\alpha(h) = \frac{1}{111 \times (16 - h)} \times \sum_{\zeta=h}^{16} \sum_{u=1}^{111} S_{\alpha, \zeta} \left[\widehat{d}_{v+\zeta}^{s, \text{lb}}(u), \widehat{d}_{v+\zeta}^{s, \text{ub}}(u), d_{v+\zeta}^s(u) \right].$$

Averaging over all forecast horizons, we obtain the overall mean interval score

$$\bar{S}_\alpha = \frac{1}{15} \sum_{h=1}^{15} \bar{S}_\alpha(h),$$

$$M[S_\alpha] = \text{median}[\bar{S}_\alpha(h)].$$

4.3. Interval forecast results

For $h = 1, 2, \dots, 15$, we present the estimated values of ξ_α obtained from the univariate functional time-series model with the CDF transformation in Table 1. Regardless of the method used to select the number of principal components retained, the values of ξ_α exhibit an increasing trend as h increases. This pattern highlights the increasing uncertainty associated with longer-term forecasts. When $h = 15$, there exists a numerical instability issue since we have only 2 years of data samples in the validation set.

For various functional time-series models with the EVR criterion to select the number of components, we evaluate and compare their ECP_h , CPD_h and $\bar{S}_{\alpha, h}$, where $h = 1, \dots, 15$. In Table 2, we

Table 2. At the nominal coverage probabilities of 80%, we evaluate and compare the interval forecast accuracy between the conformal and standard deviation approaches, measured by ECP, CPD and $\bar{S}_\alpha$, for three functional time-series models with the EVR criterion for selecting the number of components.

Sex	Metric	Approach	CDF			CLR		
			UFTS	MFTS	MLFTS	UFTS	MFTS	MLFTS
F	$\overline{\text{ECP}}$	sd	0.759	0.807	0.857	0.735	0.954	0.883
		conformal	0.694	0.729	0.776	0.594	0.852	0.809
	M[ECP]	sd	0.757	0.840	0.838	0.707	0.985	0.905
		conformal	0.671	0.741	0.769	0.601	0.870	0.804
	$\overline{\text{CPD}}$	sd	0.050	0.048	0.062	0.092	0.154	0.085
		conformal	0.106	0.071	0.037	0.206	0.055	0.047
	M[CPD]	sd	0.043	0.047	0.038	0.093	0.185	0.105
		conformal	0.129	0.059	0.032	0.199	0.070	0.047
	$\bar{S}_\alpha$	sd	470.641	450.607	279.509	440.152	304.421	282.987
		conformal	447.700	430.650	285.221	422.189	308.855	258.568
	M[$\bar{S}_\alpha$]	sd	480.484	407.819	249.753	475.842	280.564	262.005
		conformal	460.567	400.257	270.113	438.396	319.200	234.998
M	$\overline{\text{ECP}}$	sd	0.845	0.702	0.802	0.950	0.662	0.833
		conformal	0.728	0.629	0.712	0.952	0.545	0.763
	M[ECP]	sd	0.855	0.685	0.790	0.947	0.665	0.839
		conformal	0.695	0.642	0.709	0.947	0.553	0.736
	$\overline{\text{CPD}}$	sd	0.053	0.098	0.032	0.150	0.138	0.080
		conformal	0.086	0.171	0.106	0.152	0.255	0.078
	M[CPD]	sd	0.055	0.115	0.027	0.147	0.135	0.069
		conformal	0.105	0.158	0.091	0.147	0.247	0.087
	$\bar{S}_\alpha$	sd	324.628	297.966	286.152	520.602	406.046	291.875
		conformal	279.971	290.123	272.383	461.792	390.239	247.519
	M[$\bar{S}_\alpha$]	sd	341.024	309.338	296.565	548.271	433.355	299.644
		conformal	310.794	305.092	285.865	495.595	415.860	256.614

Note: Based on the ECP and CPD, we highlight in bold the functional time-series method with the smallest values for each of the two approaches.

present the averaged metrics $\overline{\text{ECP}}$, $\overline{\text{CPD}}$ and $\bar{S}_\alpha$, as well as the median $M[\text{ECP}_h]$, $M[\text{CPD}_h]$ and $M[\bar{S}_{\alpha,h}]$. At the $\alpha = 0.2$ significance level, the conformal prediction interval approach coupled with the MLFTS generally provides the smallest mean and median CPD values and interval scores for both the CDF and CLR transformations.

For female data, under the CDF transformation, the MFTS (standard deviation approach) achieves the lowest mean CPD, while the MLFTS attains the lowest median CPD and the smallest mean and median interval scores. Under the CLR transformation, the UFTS minimizes median CPD but is slightly less effective than the MLFTS in mean CPD and interval scores.

For male data, under the CDF transformation, the MLFTS achieves the lowest mean and median CPD as well as the smallest interval scores. Under the CLR transformation, the MLFTS yields the smallest mean and median CPD. Taking into account the smallest mean and median interval scores, the MLFTS is the recommended choice.

At the $\alpha = 0.05$ significance level, Table 3 highlights the conformal prediction interval with the MLFTS method as the best performer for female data under the CDF transformation. Under the CLR transformation, the MFTS method outperforms UFTS. For male data, MLFTS is recommended with the CLR transformation, while UFTS produces smaller CPD values and interval scores under the CDF transformation.

For female data, MLFTS achieves the lowest mean and median CPD, along with the smallest interval scores, under both the CDF and CLR transformations. For male data, the UFTS yields the lowest mean and median CPD, although the MLFTS and MFTS provide better interval scores. Under the CLR transformation, the UFTS attains the smallest mean and median CPD, but considering interval scores, the MLFTS remains the preferred choice. For comparison, we also consider $K = 6$ number of components and report their results in the Appendix 1.

Table 3. At the nominal coverage probabilities of 95%, we evaluate and compare the interval forecast accuracy between the conformal and standard deviation approaches, measured by ECP_h , CPD_h and $S_{\alpha,h}$, for three functional time-series models with the EVR criterion for selecting the number of components.

Sex	Metric	Approach	CDF			CLR		
			UFTS	MFTS	MLFTS	UFTS	MFTS	MLFTS
F	$\overline{ECP}$	sd	0.856	0.872	0.948	0.854	0.988	0.954
		conformal	0.761	0.795	0.869	0.688	0.909	0.886
	M[ECP]	sd	0.868	0.885	0.958	0.857	0.999	0.962
		conformal	0.751	0.795	0.887	0.715	0.912	0.905
	$\overline{CPD}$	sd	0.094	0.078	0.023	0.096	0.038	0.023
		conformal	0.189	0.155	0.081	0.262	0.041	0.064
	M[CPD]	sd	0.082	0.065	0.023	0.093	0.049	0.023
		conformal	0.199	0.155	0.063	0.235	0.038	0.045
	$\overline{S}_\alpha$	sd	902.186	834.717	455.273	744.670	453.627	452.517
		conformal	1068.769	922.037	410.163	846.837	346.150	335.551
	M[S_α]	sd	926.920	763.808	324.887	702.110	407.010	368.169
		conformal	1012.438	795.885	370.749	841.039	342.580	358.339
M	$\overline{ECP}$	sd	0.946	0.886	0.921	0.977	0.873	0.939
		conformal	0.798	0.718	0.784	0.975	0.636	0.853
	M[ECP]	sd	0.953	0.873	0.908	0.985	0.867	0.929
		conformal	0.743	0.730	0.755	0.980	0.637	0.851
	$\overline{CPD}$	sd	0.018	0.068	0.040	0.031	0.078	0.038
		conformal	0.152	0.232	0.167	0.026	0.314	0.097
	M[CPD]	sd	0.014	0.077	0.042	0.035	0.083	0.036
		conformal	0.207	0.220	0.195	0.030	0.313	0.099
	$\overline{S}_\alpha$	sd	493.131	475.002	456.609	660.490	611.946	460.363
		conformal	422.177	579.461	501.265	522.754	839.853	368.778
	M[S_α]	sd	448.888	425.326	445.072	750.222	552.793	447.374
		conformal	457.082	572.234	483.798	579.485	848.386	374.045

5. Conclusion

We propose a general strategy for constructing prediction intervals for the age distribution of death. This approach leverages a validation set to determine an optimal tuning parameter that aligns empirical and nominal coverage probabilities. Using this optimized parameter, we construct prediction intervals for the testing set.

To illustrate the effectiveness of this strategy, we analyze Japanese age- and sex-specific life-table death counts, comparing three functional time-series forecasting models: univariate, multivariate, and multilevel functional time-series models. Our findings suggest that the multilevel functional time-series method generally performs best. Additionally, when selecting the number of components, we find little difference between the EVR criterion and the setting $K = 6$. Given that overfitting does not adversely affect the accuracy of the forecast, we recommend the latter.

Using the age distribution of deaths, we present our methodology for constructing distribution-free and model-agnostic prediction intervals. Other measures of mortality, such as age-specific mortality rates or hazard rates, could also be considered in the modeling. In Appendices 2 and 3, we demonstrate our proposed sd approach for constructing prediction intervals and evaluating its empirical coverage probability using the Australian and Canadian age- and sex-specific log mortality rates, respectively. Using the proposed sd approach, it achieves superior finite-sample coverage probability in comparison to the classical functional time-series model of Hyndman and Ullah (2007). The method of Hyndman and Ullah (2007) computes the total variance and constructs the prediction interval parametrically based on the assumption of a Gaussian distribution.

There are at least five ways in which the methodology can be extended. (1) The functional standard deviation was computed coordinate-wise. Several functional depths exist, which can be implemented to compute other variants of standard deviations. (2) Instead of symmetric prediction intervals, one can consider asymmetric ones. In that case, two tuning parameters are needed to adjust the lower

and upper bounds. (3) The data set was divided equally into training, validation, and testing samples. Other proportions may be possible and lead to a more accurate selection of the tuning parameter ξ_α and more accurate interval forecasts. (4) For demonstration, we implemented a suite of functional time-series models. Other time-series extrapolation models may also be considered. (5) We use the life-table data directly in our modeling, but we could extend the analysis by incorporating their estimation error into the model, reflecting the underlying observational data used to construct the life table.

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Appendices

Appendix 1. Interval forecast results when $K = 6$

In Tables A1 and A2, we present the interval forecast accuracy between the conformal and standard deviation approaches for three functional time-series models with the first six retained components at the nominal coverage probabilities of 80% and 95%, respectively.

From Table A1, the MFTS and MLFTS outperform the UFTS with smaller mean and median CPD and interval scores for both approaches. However, for male data, the UFTS achieves a lower CPD using the standard deviation approach with the CLR transformation.

From Table A2, the MFTS and MLFTS outperform the UFTS with smaller mean and median CPD and interval scores for both approaches. However, for male data, the UFTS achieves a lower CPD and an interval score using the standard deviation approach with the CDF transformation.

Appendix 2. Australian age-specific mortality rates

We analyze Australian age- and sex-specific mortality rates spanning from 1921 to 2020, obtained from Human Mortality Database (2025). These rates represent the ratio of death counts to population exposure in each respective year and age group (based on one-year intervals). Our study covers age groups from 0 to 99 in single years, with the final group covering ages 100 and above. Age-specific mortality rates are often modeled and forecasted using natural logarithmic transformations. In Figure A1, we present rainbow plots for log mortality rates.

In Table A3, we compute the empirical coverage probability specific to each horizon and its coverage probability difference for the sd approach at the nominal coverage probability 80%. For comparison, we implement the parametric

Table A1. At the nominal coverage probabilities of 80%, we evaluate and compare the interval forecast accuracy between the conformal and standard deviation approaches, measured by ECP, CPD and S_α , for three functional time-series models with $K = 6$.

Sex	Metric	Approach	CDF			CLR		
			UFTS	MFTS	MLFTS	UFTS	MFTS	MLFTS
F	$\overline{\text{ECP}}$	sd	0.750	0.801	0.833	0.785	0.835	0.878
		conformal	0.681	0.726	0.766	0.654	0.815	0.753
	M[ECP]	sd	0.746	0.809	0.832	0.771	0.821	0.873
		conformal	0.662	0.726	0.764	0.677	0.825	0.755
	$\overline{\text{CPD}}$	sd	0.061	0.036	0.037	0.050	0.071	0.083
		conformal	0.119	0.074	0.035	0.146	0.065	0.061
	M[CPD]	sd	0.054	0.035	0.034	0.042	0.066	0.073
		conformal	0.138	0.074	0.036	0.123	0.072	0.045
	$\overline{S}_\alpha$	sd	425.527	336.788	309.193	376.570	309.369	274.407
		conformal	404.812	327.194	311.363	351.138	256.756	259.693
	M[S_α]	sd	422.625	318.620	264.745	362.494	281.715	232.967
		conformal	408.239	324.021	293.286	367.602	265.018	270.223
M	$\overline{\text{ECP}}$	sd	0.815	0.737	0.825	0.854	0.747	0.887
		conformal	0.701	0.652	0.740	0.768	0.652	0.810
	M[ECP]	sd	0.820	0.735	0.820	0.841	0.733	0.875
		conformal	0.669	0.655	0.728	0.783	0.646	0.819
	$\overline{\text{CPD}}$	sd	0.035	0.063	0.026	0.057	0.058	0.087
		conformal	0.101	0.148	0.068	0.066	0.148	0.025
	M[CPD]	sd	0.034	0.065	0.020	0.041	0.067	0.075
		conformal	0.131	0.145	0.072	0.049	0.154	0.021
	$\overline{S}_\alpha$	sd	332.674	318.413	289.418	336.100	444.925	275.450
		conformal	285.552	302.708	263.287	279.133	406.393	234.783
	M[S_α]	sd	354.513	347.581	331.476	338.271	457.361	289.172
		conformal	327.190	308.440	298.650	274.904	448.065	257.413

Note: Based on the ECP and CPD, we highlight in bold the functional time-series method with the smallest values for each of the two approaches.

Table A2. At the nominal coverage probabilities of 95%, we evaluate and compare the interval forecast accuracy between the conformal and standard deviation approaches, measured by ECP_h , CPD_h and $S_{\alpha,h}$, for three functional time-series models with $K = 6$.

Sex	Metric	Approach	CDF			CLR		
			UFTS	MFTS	MLFTS	UFTS	MFTS	MLFTS
F	$\overline{ECP}$	sd	0.859	0.891	0.931	0.900	0.924	0.963
		conformal	0.755	0.801	0.839	0.745	0.887	0.833
	M[ECP]	sd	0.865	0.884	0.929	0.892	0.923	0.971
		conformal	0.742	0.806	0.824	0.776	0.868	0.841
	$\overline{CPD}$	sd	0.091	0.059	0.023	0.050	0.048	0.022
		conformal	0.195	0.149	0.111	0.205	0.064	0.117
	M[CPD]	sd	0.085	0.066	0.023	0.058	0.042	0.021
		conformal	0.208	0.144	0.126	0.174	0.082	0.109
	$\bar{S}_\alpha$	sd	783.946	582.336	491.354	613.577	557.077	434.461
		conformal	943.683	652.737	574.971	606.063	379.501	341.970
	M[S_α]	sd	732.045	510.697	371.856	532.389	445.167	326.626
		conformal	852.400	573.037	437.359	521.617	343.097	358.741
M	$\overline{ECP}$	sd	0.946	0.882	0.937	0.942	0.867	0.961
		conformal	0.771	0.742	0.814	0.848	0.749	0.877
	M[ECP]	sd	0.953	0.872	0.938	0.945	0.864	0.956
		conformal	0.717	0.744	0.786	0.851	0.747	0.886
	$\overline{CPD}$	sd	0.018	0.068	0.022	0.030	0.083	0.015
		conformal	0.179	0.208	0.136	0.102	0.201	0.073
	M[CPD]	sd	0.014	0.078	0.019	0.027	0.086	0.008
		conformal	0.233	0.206	0.164	0.099	0.203	0.064
	$\bar{S}_\alpha$	sd	493.131	602.081	493.277	662.022	762.599	462.813
		conformal	454.399	632.817	418.730	442.703	874.558	327.106
	M[S_α]	sd	448.888	562.958	477.297	575.785	836.510	408.421
		conformal	518.473	639.114	434.471	354.886	886.527	333.583

approach of Hyndman and Ullah (2007) implemented in the `ftsa` package in . Using data from 1921 to 1976, we computed the forecasts for the validation period from 1977 to 1998. For $h = 1, 2, \dots, 21$, we determine the optimal tuning parameters, with which we evaluate the empirical coverage probability based on the test period from 1999 to 2020. The proposed sd approach achieves superior finite-sample coverage in comparison to the parametric approach based on the total variance under the Gaussian distribution assumption.

In Table A4, we report the empirical coverage probability specific to each horizon and its coverage probability difference for the sd approach at the nominal level 95%. Compared with the parametric approach of Hyndman and Ullah (2007), the sd approach achieves superior finite-sample coverage probability and often produces empirical coverage probability at and above the nominal level.

Appendix 3. Canadian age-specific mortality rates

We also analyze Canadian age- and sex-specific mortality rates spanning from 1921 to 2022, obtained from the Human Mortality Database (2025). Our study covers age groups from 0 to 99 in single years, with the final group covering ages 100 and above. We present rainbow plots for log mortality rates in Figure A2, where the data from the distant past are shown in red and the more recent data in purple.

In Table A5, we compute the horizon-specific empirical coverage probability and its coverage probability difference for the sd approach at the 80% nominal coverage probability. Using data from 1921 to 1978, we computed the forecasts for the validation period from 1979 to 2000. For $h = 1, 2, \dots, 21$, we determine the optimal tuning parameters, with which we evaluate the empirical coverage probability based on the test period from 2001 to 2022. The sd approach achieves superior finite-sample coverage probability and often produces empirical coverage probability around the nominal level.

In Table A6, we display the horizon-specific empirical coverage probability and its coverage probability difference for the sd approach at the 95% nominal coverage probability. Compared with the parametric approach of Hyndman and Ullah (2007), the sd approach achieves superior finite-sample coverage probability and often produces empirical coverage probability at and above the nominal level.

Table A3. At the nominal coverage probability of 80%, we compute the empirical coverage probability and its coverage probability difference between the sd approach and parametric approach.

<i>h</i>	sd approach								parametric approach							
	Female				Male				Female				Male			
	Smooth		Raw		Smooth		Raw		Smooth		Raw		Smooth		Raw	
	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD
1	0.807	0.007	0.802	0.002	0.815	0.015	0.849	0.049	0.545	0.255	0.746	0.054	0.573	0.227	0.738	0.062
2	0.819	0.019	0.836	0.036	0.846	0.046	0.849	0.049	0.544	0.256	0.749	0.051	0.529	0.271	0.686	0.114
3	0.815	0.015	0.838	0.038	0.875	0.075	0.875	0.075	0.552	0.248	0.739	0.061	0.498	0.302	0.645	0.155
4	0.829	0.029	0.845	0.045	0.903	0.103	0.900	0.100	0.553	0.247	0.730	0.070	0.476	0.324	0.594	0.206
5	0.824	0.024	0.842	0.042	0.923	0.123	0.926	0.126	0.558	0.242	0.726	0.074	0.450	0.350	0.563	0.237
6	0.832	0.032	0.856	0.056	0.949	0.149	0.945	0.145	0.562	0.238	0.715	0.085	0.425	0.375	0.533	0.267
7	0.822	0.022	0.853	0.053	0.954	0.154	0.957	0.157	0.544	0.256	0.697	0.103	0.405	0.395	0.502	0.298
8	0.825	0.025	0.861	0.061	0.980	0.180	0.983	0.183	0.547	0.253	0.700	0.100	0.389	0.411	0.496	0.304
9	0.822	0.022	0.873	0.073	0.993	0.193	0.988	0.188	0.557	0.243	0.690	0.110	0.380	0.420	0.497	0.303
10	0.819	0.019	0.881	0.081	0.989	0.189	0.984	0.184	0.560	0.240	0.691	0.109	0.369	0.431	0.486	0.314
11	0.830	0.030	0.870	0.070	0.979	0.179	0.979	0.179	0.547	0.253	0.676	0.124	0.371	0.429	0.481	0.319
12	0.833	0.033	0.887	0.087	0.977	0.177	0.984	0.184	0.557	0.243	0.671	0.129	0.364	0.436	0.466	0.334
13	0.849	0.049	0.885	0.085	0.983	0.183	0.990	0.190	0.566	0.234	0.673	0.127	0.366	0.434	0.471	0.329
14	0.827	0.027	0.914	0.114	0.983	0.183	0.985	0.185	0.554	0.246	0.673	0.127	0.372	0.428	0.483	0.317
15	0.830	0.030	0.901	0.101	0.978	0.178	0.978	0.178	0.561	0.239	0.663	0.137	0.395	0.405	0.459	0.341
16	0.789	0.011	0.881	0.081	0.969	0.169	0.963	0.163	0.549	0.251	0.652	0.148	0.402	0.398	0.463	0.337
17	0.807	0.007	0.875	0.075	0.969	0.169	0.959	0.159	0.540	0.260	0.645	0.155	0.408	0.392	0.464	0.336
18	0.794	0.006	0.861	0.061	0.970	0.170	0.956	0.156	0.507	0.293	0.650	0.150	0.420	0.380	0.461	0.339
19	0.829	0.029	0.869	0.069	0.978	0.178	0.975	0.175	0.505	0.295	0.641	0.159	0.426	0.374	0.475	0.325
20	0.825	0.025	0.931	0.131	0.927	0.127	0.921	0.121	0.492	0.308	0.647	0.153	0.419	0.381	0.488	0.312
21	0.856	0.056	0.931	0.131	0.876	0.076	0.921	0.121	0.550	0.250	0.673	0.127	0.426	0.374	0.485	0.315
Mean	0.823	0.025	0.871	0.071	0.944	0.144	0.946	0.146	0.545	0.255	0.688	0.112	0.422	0.378	0.521	0.279

Table A4. At the nominal coverage probability of 95%, we compute the empirical coverage probability and its coverage probability difference between the sd approach and parametric approach.

<i>h</i>	sd approach								parametric approach							
	Female				Male				Female				Male			
	Smooth		Raw		Smooth		Raw		Smooth		Raw		Smooth		Raw	
	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD
1	0.933	0.017	0.938	0.012	0.949	0.001	0.964	0.014	0.715	0.235	0.907	0.043	0.752	0.198	0.899	0.051
2	0.938	0.012	0.943	0.007	0.972	0.022	0.980	0.030	0.722	0.228	0.906	0.044	0.736	0.214	0.883	0.067
3	0.943	0.007	0.952	0.002	0.973	0.023	0.982	0.032	0.722	0.228	0.905	0.045	0.707	0.243	0.848	0.102
4	0.947	0.003	0.957	0.007	0.981	0.031	0.984	0.034	0.741	0.209	0.899	0.051	0.688	0.262	0.826	0.124
5	0.952	0.002	0.964	0.014	0.989	0.039	0.988	0.038	0.737	0.213	0.893	0.057	0.644	0.306	0.806	0.144
6	0.944	0.006	0.956	0.006	0.991	0.041	0.997	0.047	0.732	0.218	0.888	0.062	0.620	0.330	0.775	0.175
7	0.940	0.010	0.958	0.008	0.993	0.043	0.991	0.041	0.727	0.223	0.876	0.074	0.592	0.358	0.751	0.199
8	0.947	0.003	0.964	0.014	0.999	0.049	0.997	0.047	0.723	0.227	0.884	0.066	0.585	0.365	0.739	0.211
9	0.937	0.013	0.964	0.014	1.000	0.050	1.000	0.050	0.723	0.227	0.883	0.067	0.574	0.376	0.732	0.218
10	0.941	0.009	0.966	0.016	1.000	0.050	1.000	0.050	0.717	0.233	0.858	0.092	0.561	0.389	0.720	0.230
11	0.947	0.003	0.972	0.022	0.998	0.048	0.993	0.043	0.715	0.235	0.861	0.089	0.568	0.382	0.705	0.245
12	0.945	0.005	0.977	0.027	0.995	0.045	1.000	0.050	0.698	0.252	0.849	0.101	0.557	0.393	0.697	0.253
13	0.944	0.006	0.977	0.027	0.995	0.045	0.999	0.049	0.691	0.259	0.846	0.104	0.574	0.376	0.693	0.257
14	0.944	0.006	0.981	0.031	0.994	0.044	0.994	0.044	0.691	0.259	0.838	0.112	0.582	0.368	0.691	0.259
15	0.965	0.015	0.984	0.034	0.996	0.046	0.990	0.040	0.689	0.261	0.832	0.118	0.571	0.379	0.676	0.274
16	0.949	0.001	0.987	0.037	0.992	0.042	0.982	0.032	0.679	0.271	0.826	0.124	0.576	0.374	0.668	0.282
17	0.957	0.007	0.975	0.025	0.995	0.045	0.990	0.040	0.685	0.265	0.820	0.130	0.576	0.374	0.660	0.290
18	0.937	0.013	0.978	0.028	0.996	0.046	0.980	0.030	0.685	0.265	0.822	0.128	0.560	0.390	0.659	0.291
19	0.955	0.005	0.980	0.030	1.000	0.050	1.000	0.050	0.671	0.279	0.832	0.118	0.554	0.396	0.661	0.289
20	0.944	0.006	0.990	0.040	0.990	0.040	0.990	0.040	0.677	0.273	0.805	0.145	0.548	0.402	0.644	0.306
21	0.975	0.025	0.970	0.020	0.965	0.015	0.990	0.040	0.723	0.227	0.837	0.113	0.554	0.396	0.673	0.277
Mean	0.947	0.008	0.968	0.020	0.989	0.039	0.990	0.040	0.708	0.242	0.860	0.090	0.604	0.346	0.734	0.216

Table A5. At the nominal coverage probability of 80%, we compute the empirical coverage probability and its coverage probability difference between the sd approach and parametric approach for the Canadian data.

<i>h</i>	sd approach								parametric approach							
	Female				Male				Female				Male			
	Smooth		Raw		Smooth		Raw		Smooth		Raw		Smooth		Raw	
	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD
1	0.769	0.031	0.775	0.025	0.742	0.058	0.728	0.072	0.571	0.229	0.766	0.034	0.498	0.302	0.649	0.151
2	0.773	0.027	0.786	0.014	0.738	0.062	0.712	0.088	0.588	0.212	0.761	0.039	0.444	0.356	0.587	0.213
3	0.768	0.032	0.764	0.036	0.726	0.074	0.702	0.098	0.566	0.234	0.732	0.068	0.403	0.397	0.535	0.265
4	0.770	0.030	0.755	0.045	0.736	0.064	0.714	0.086	0.581	0.219	0.736	0.064	0.398	0.402	0.508	0.292
5	0.743	0.057	0.733	0.067	0.734	0.066	0.724	0.076	0.557	0.243	0.691	0.109	0.372	0.428	0.481	0.319
6	0.744	0.056	0.732	0.068	0.740	0.060	0.726	0.074	0.543	0.257	0.680	0.120	0.362	0.438	0.467	0.333
7	0.749	0.051	0.731	0.069	0.749	0.051	0.729	0.071	0.545	0.255	0.666	0.134	0.353	0.447	0.452	0.348
8	0.745	0.055	0.716	0.084	0.760	0.040	0.737	0.063	0.530	0.270	0.647	0.153	0.341	0.459	0.439	0.361
9	0.744	0.056	0.732	0.068	0.793	0.007	0.778	0.022	0.521	0.279	0.632	0.168	0.345	0.455	0.438	0.362
10	0.724	0.076	0.706	0.094	0.815	0.015	0.784	0.016	0.515	0.285	0.618	0.182	0.350	0.450	0.427	0.373
11	0.723	0.077	0.713	0.087	0.836	0.036	0.800	0.000	0.505	0.295	0.601	0.199	0.366	0.434	0.441	0.359
12	0.734	0.066	0.714	0.086	0.883	0.083	0.846	0.046	0.485	0.315	0.585	0.215	0.351	0.449	0.431	0.369
13	0.747	0.053	0.730	0.070	0.850	0.050	0.828	0.028	0.503	0.297	0.589	0.211	0.356	0.444	0.434	0.366
14	0.740	0.060	0.722	0.078	0.862	0.062	0.810	0.010	0.492	0.308	0.585	0.215	0.366	0.434	0.430	0.370
15	0.751	0.049	0.713	0.087	0.829	0.029	0.797	0.003	0.460	0.340	0.546	0.254	0.351	0.449	0.415	0.385
16	0.692	0.108	0.683	0.117	0.851	0.051	0.843	0.043	0.429	0.371	0.506	0.294	0.341	0.459	0.407	0.393
17	0.705	0.095	0.691	0.109	0.855	0.055	0.835	0.035	0.422	0.378	0.490	0.310	0.337	0.463	0.381	0.419
18	0.713	0.087	0.695	0.105	0.899	0.099	0.877	0.077	0.414	0.386	0.497	0.303	0.327	0.473	0.386	0.414
19	0.782	0.018	0.755	0.045	0.943	0.143	0.928	0.128	0.411	0.389	0.520	0.280	0.329	0.471	0.351	0.449
20	0.746	0.054	0.634	0.166	0.881	0.081	0.878	0.078	0.360	0.440	0.469	0.331	0.314	0.486	0.333	0.467
21	0.762	0.038	0.634	0.166	0.792	0.008	0.832	0.032	0.287	0.513	0.391	0.409	0.337	0.463	0.337	0.463
Mean	0.744	0.056	0.720	0.080	0.810	0.057	0.791	0.055	0.490	0.310	0.605	0.195	0.364	0.436	0.444	0.356

Table A6. At the nominal coverage probability of 95%, we compute the empirical coverage probability and its coverage probability difference between the sd approach and parametric approach.

<i>h</i>	sd approach								parametric approach							
	Female				Male				Female				Male			
	Smooth		Raw		Smooth		Raw		Smooth		Raw		Smooth		Raw	
	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD	ECP	CPD
1	0.918	0.032	0.907	0.043	0.892	0.058	0.881	0.069	0.738	0.212	0.902	0.048	0.683	0.267	0.845	0.105
2	0.925	0.025	0.915	0.035	0.876	0.074	0.864	0.086	0.746	0.204	0.897	0.053	0.613	0.337	0.778	0.172
3	0.906	0.044	0.900	0.050	0.855	0.095	0.851	0.099	0.733	0.217	0.865	0.085	0.560	0.390	0.710	0.240
4	0.916	0.034	0.903	0.047	0.855	0.095	0.858	0.092	0.735	0.215	0.865	0.085	0.538	0.412	0.670	0.280
5	0.895	0.055	0.881	0.069	0.864	0.086	0.848	0.102	0.704	0.246	0.837	0.113	0.528	0.422	0.642	0.308
6	0.878	0.072	0.867	0.083	0.871	0.079	0.858	0.092	0.694	0.256	0.822	0.128	0.497	0.453	0.615	0.335
7	0.877	0.073	0.860	0.090	0.876	0.074	0.858	0.092	0.691	0.259	0.806	0.144	0.494	0.456	0.612	0.338
8	0.869	0.081	0.844	0.106	0.894	0.056	0.878	0.072	0.673	0.277	0.788	0.162	0.498	0.452	0.601	0.349
9	0.866	0.084	0.854	0.096	0.895	0.055	0.875	0.075	0.656	0.294	0.778	0.172	0.487	0.463	0.588	0.362
10	0.847	0.103	0.840	0.110	0.902	0.048	0.882	0.068	0.648	0.302	0.750	0.200	0.489	0.461	0.586	0.364
11	0.851	0.099	0.847	0.103	0.920	0.030	0.903	0.047	0.634	0.316	0.737	0.213	0.506	0.444	0.593	0.357
12	0.869	0.081	0.829	0.121	0.954	0.004	0.938	0.012	0.628	0.322	0.714	0.236	0.515	0.435	0.598	0.352
13	0.869	0.081	0.847	0.103	0.935	0.015	0.940	0.010	0.630	0.320	0.727	0.232	0.525	0.425	0.592	0.358
14	0.834	0.116	0.826	0.124	0.964	0.014	0.978	0.028	0.617	0.333	0.722	0.228	0.528	0.422	0.601	0.349
15	0.838	0.112	0.832	0.118	0.944	0.006	0.978	0.028	0.597	0.353	0.712	0.238	0.519	0.431	0.580	0.370
16	0.825	0.125	0.818	0.132	0.963	0.013	0.972	0.022	0.562	0.388	0.673	0.277	0.506	0.444	0.567	0.383
17	0.810	0.140	0.784	0.166	0.972	0.022	0.979	0.029	0.558	0.392	0.649	0.301	0.493	0.457	0.569	0.381
18	0.808	0.142	0.810	0.140	0.980	0.030	0.974	0.024	0.562	0.388	0.657	0.293	0.497	0.453	0.543	0.407
19	0.879	0.071	0.901	0.049	0.998	0.048	0.978	0.028	0.564	0.386	0.653	0.297	0.460	0.490	0.532	0.418
20	0.911	0.039	0.908	0.042	0.990	0.040	0.970	0.020	0.521	0.429	0.640	0.310	0.426	0.524	0.475	0.475
21	0.881	0.069	0.876	0.074	0.980	0.030	0.941	0.009	0.480	0.470	0.634	0.316	0.431	0.519	0.480	0.470
Mean	0.870	0.080	0.859	0.091	0.923	0.046	0.914	0.053	0.637	0.313	0.754	0.196	0.514	0.436	0.608	0.342

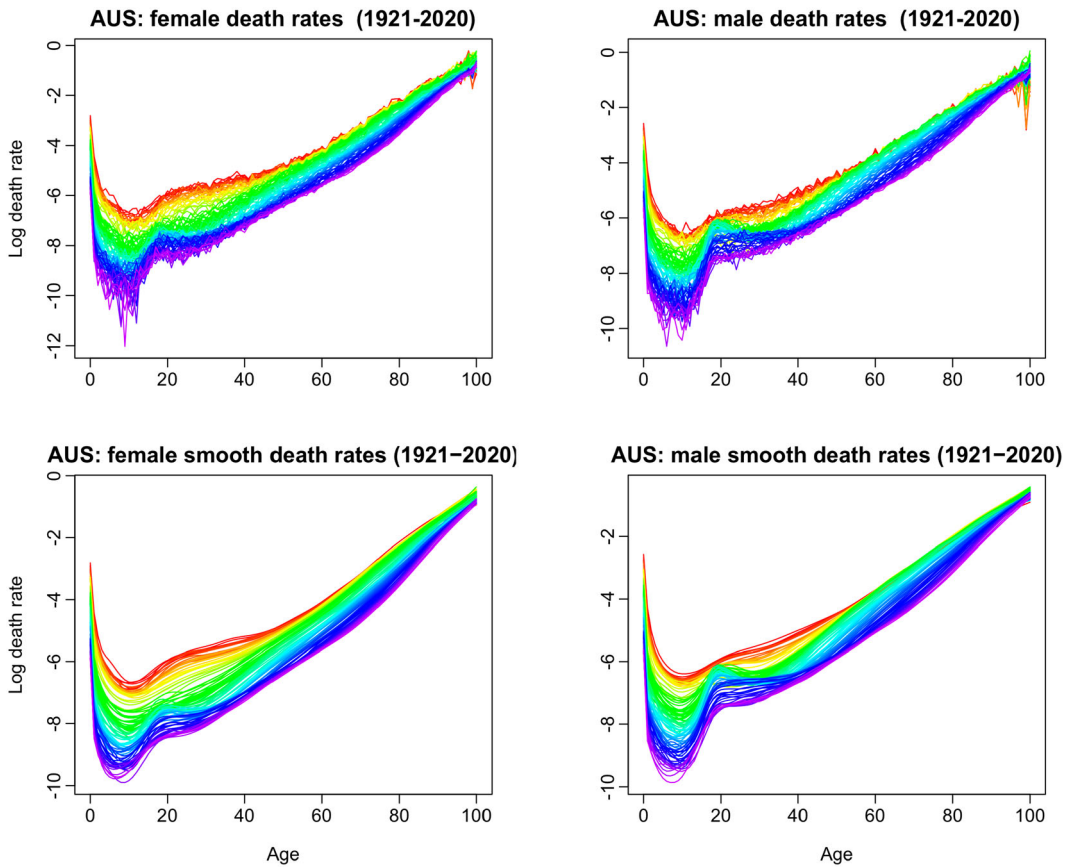


Figure A1. Rainbow plots of the original and smoothed age-specific mortality rates for the Australian female and male data from 1921 to 2020. Smoothing was performed via penalized spline with monotonic constraint described in Hyndman and Ullah (2007).

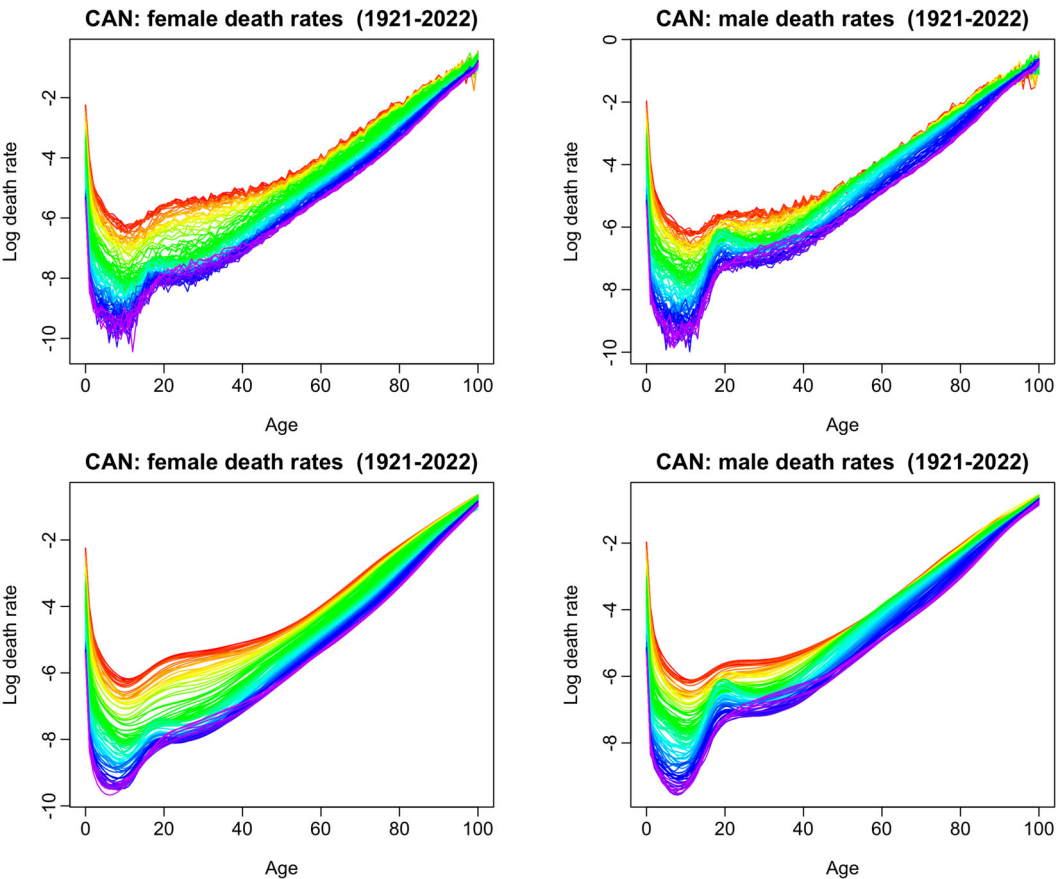


Figure A2. Rainbow plots of the original and smoothed age-specific mortality rates for the Canadian female and male data from 1921 to 2022.