"Historical Studies on Mortality," special issue, *Canadian Studies in Population* 39, no. 3–4 (2012): 79–106.

Regional disparities in Canadian adult and old-age mortality: A comparative study based on smoothed mortality ratio surfaces and age at death distributions

Nadine Ouellette
Department of Demography, University of California, Berkeley
nadineo@demog.berkeley.edu

Robert Bourbeau
Département de Démographie, Université de Montréal

Carlo G. Camarda
Institut national d'études démographiques (INED), Paris

Abstract

This paper examines adult and old-age mortality differentials in Canada between 1930 and 2007 at the provincial level, using the Canadian Human Mortality Database and the flexible smoothing P-spline method in two-dimensions well-suited to the study of small populations. Our analysis reveals that provincial disparities in adult mortality in general, and among the elderly population in particular, are substantial in Canada. Moreover, based on the modal age at death and the standard deviation of ages at death above the mode, provincial disparities at older ages have barely reduced over time, despite the great mortality improvements in all provinces since the early 20th century. In the last few years studied, evidence of the shifting mortality regime was found among females in most Western and Central provinces, while all males were still undergoing an old-age mortality compression regime.

Keywords: *adult mortality, modal age at death, old-age mortality compression, Canadian provinces, P-spline smoothing.*

Résumé

Cet article porte sur les disparités provinciales en matière de mortalité chez les adultes et les personnes âgées au Canada entre 1930 et 2007. Les données, provenant de la Base de données sur la longévité canadienne, sont analysées par le biais d'une méthode de lissage par P-splines en deux dimensions bien adaptée à l'étude des populations de petite taille. Notre étude révèle que les écarts de mortalité entre les provinces, chez les adultes en général et chez les personnes âgées en particulier, sont importants au Canada. Sur la base des trajectoires temporelles de l'âge modal au décès et de l'écart type des âges au décès situés au-delà du mode, les disparités provinciales aux grands âges ont à peine diminué. Cela malgré la baisse notable de la mortalité dans toutes les provinces depuis le début du 20e siècle. Enfin, au cours des dernières années, le scénario de compression de la mortalité aux grands âges était toujours en vigueur chez les hommes. Chez les femmes, dans la plupart des provinces de l'Ouest et du centre du pays, nous observons plutôt un déplacement de l'ensemble des durées de vie adultes vers des âges plus élevés, sans réduction parallèle de la dispersion de la mortalité aux grands âges.

Mots-clés : *mortalité adulte, âge modal au décès, compression de la mortalité aux grands âges, provinces canadiennes, lissage par P-splines.*

Introduction

As in many low-mortality countries around the globe, the 20th century has been an important milestone in the evolution of epidemiological conditions in Canada. With the fall in infant mortality, infectious and parasitic diseases, and (more recently) cardiovascular diseases and cancer, life expectancy at birth rose considerably during this time.

Indeed, total life expectancy at birth in Canada was about 47 years in 1901 (Bourbeau et al. 1997) and reached almost 81 years in 2007 (CHMD 2010), yielding a spectacular increase of more than 30 years in about a century. Over this period, Canadian females have systematically enjoyed higher life expectancy at birth than males. This long-lasting female advantage is likely to be due to a complex interplay of many biological, social, and behavioral factors (Madigan 1957; Waldron 1976; Lopez and Ruzicka 1983; Vallin 1993). However, the sex differential in life expectancy peaked at about 7 years in the 1970s in Canada, and has been narrowing ever since (Trovato and Lalu 1995, 2001; Nault 1997; Manuel and Hockin 2000; Zanfongnon 2008). In recent decades, males have indeed been more successful than females in reducing their mortality from cardiovascular disease and external causes of death such as accidents, suicide, and violence. Furthermore, female lung cancer mortality has increased, while decreasing among males. According to the latest data available, the sex gap in life expectancy shrank to 4.6 years in 2007 (CHMD 2010).

In terms of regional mortality conditions in Canada, all provinces and territories have recorded substantial mortality progress since the beginning of the 20th century. Previous studies have shown that these mortality improvements tended to reduce any disparities that prevailed among them (Field 1980; Wilkins 1980; Adams 1990; Nault 1997; Manuel and Hockin 2000; Prud'homme 2007). Such a finding is in line with the proposal that Canadian regions are becoming more and more homogeneous in terms of their economic and socio-demographic characteristics over time (Matthews and Davis 1986; Goyder 1993). In spite of that, regional mortality differentials persist in Canada. For example, long-standing geographical disparities in favour of provinces in the Western part of Canada compared to those in the East continue to exist. Recent contributions (Manuel and Hockin 2000; Prud'homme 2007) even reveal that since the end of the 1980s, the general east–west gradient might have clarified further into a well-defined east–central–prairies–west gradient, thus confirming that regional mortality disparities are still worth documenting in Canada.

The fact that Canada consists of ten provinces and three territories that remain disparate geographically, economically, politically, culturally, and socially undoubtedly explains part of the regional mortality differentials still recorded in the country (Trovato and Lalu 2001; Prud'homme 2007). Studies reporting the prevalence of several major well-established risk behaviors and health conditions for cardiovascular disease by province help elucidate these differentials further (Health Canada 1995; PHAC 2009). For example, the life expectancy disadvantage of the easternmost Canadian provinces—namely, the Atlantic provinces comprising New Brunswick, Nova Scotia, Prince Edward Island, and Newfoundland and Labrador—with respect to the other provinces is clearly reflected in their higher prevalence of smoking, physical inactivity, high blood pressure, and obesity. In contrast, British Columbia, Canada's westernmost province, almost systematically displays the lowest prevalence of these indicators.

Nativity composition of the population in each province also proves helpful for understanding regional mortality differentials in Canada. Indeed, previous work based on Canadian data has indicated that immigrants generally tend to experience lower levels of mortality compared to the Canadian-born population (Sharma et al. 1990; Trovato 1993; Chen et al. 1996; Bourbeau 2002). Thus, the fact that immigrants account for a much larger share of the total population in Western provinces, especially British Columbia, than in Atlantic provinces is likely to favour the former ones in terms of overall mortality conditions. Furthermore, immigrants in Western provinces may even experience lower mortality than those in Atlantic provinces because a greater part of them originate from non-European countries, specifically Asia (Chen et al. 1996; Bourbeau 2002).[1]

When the probability of surviving to older ages increases over time among individuals in a given region or within human populations in general, life expectancy at birth rises, and the shape of the survival curve usually progressively becomes more rectangular. This phenomenon, well-known as the *rectangularization of the survival curve*, is associated with a reduction in the variability of age at death (Wilmoth and Horiuchi 1999), commonly referred to as *compression of mortality*. Indeed, when compression of mortality is at work, deaths are concentrated into a shorter age interval over time, the dispersion of the age at death distribution is reduced, and the downward slope of the survival curve becomes steeper, resulting in a more rectangular shape. An abundant body of demographic and epidemiological literature has already focused on the topics of rectangularization of the survival curve and compression of mortality, providing evidence of these phenomena in Canada and several low-mortality countries and regions (Fries 1980; My-

1. Chen et al. (1996) and Bourbeau (2002) studied mortality differentials by nativity in Canada and considered three categories with respect to place of birth: the *Canadian-born* category obviously included those born in Canada, the *European immigrants* are those born in Europe, Australia, or New Zealand, and the *non-European immigrants* are those born in the remaining countries. The authors show that immigrants, especially those from non-European countries, have lower mortality rates than the Canadian-born at all ages.

ers and Manton 1984a, 1984b; Nagnur 1986; Manton and Tolley 1991; Hill 1993; Eakin and Witten 1995; Nusselder and Mackenbach 1996, 1997; Paccaud et al. 1998; Wilmoth and Horiuchi 1999; Kannisto 2000, 2001, 2007; Lynch and Brown 2001; Robine 2001; Martel 2002; Martel and Bourbeau 2003; Cheung et al. 2005, 2008, 2009; Cheung and Robine 2007; Thatcher et al. 2010; Ouellette and Bourbeau 2011).

These studies also suggest that roughly up until the 1950s, important mortality reductions among infants, children, and even young adults led to a strong compression of the overall age at death distribution in the various countries and regions over time. Afterwards, however, this overall compression of mortality slowed down substantially even though important mortality gains started to be recorded consistently among older adults (Kannisto et al. 1994; Jeune and Vaupel 1995). As pointed out by Thatcher et al. (2010), these findings have encouraged researchers to distinguish *old-age* mortality compression from *overall* mortality compression, and to study changes in the age at death distribution at *older ages* and over the *entire age range* separately. Moreover, the *late life modal age at death*[2] (referred to hereafter as *modal age at death*) and the variability of deaths around this modal age, often measured by the *standard deviation of individual life durations above the modal age at death*, emerged as an important set of tools to summarize and monitor changes in the age at death distribution at older ages over time (Kannisto 2000, 2001, 2007; Cheung et al. 2005, 2008, 2009; Cheung and Robine 2007; Canudas-Romo 2008; Thatcher et al. 2010; Ouellette and Bourbeau 2011). Indeed, unlike life expectancy at birth, which is highly sensitive to changes in mortality among infants and children, the modal age at death is solely influenced by adult and old-age mortality, and consequently much more sensitive to changes occurring among the elderly population (Kannisto 2001; Horiuchi 2003; Cheung and Robine 2007; Canudas-Romo 2010).

Rectangularization of the survival curve and compression of mortality in Canada has already been explored in a few studies. Indeed, the work by Nagnur (1986) focuses on the Canadian mortality experience from 1921 to 1981 and reveals that rectangularization of the survival curve did occur in this country during these six decades. Martel (2002) and Martel and Bourbeau (2003) take a close look at the situation in the province of Quebec between 1921 and 2000. They demonstrate that rectangularization of the survival curve and overall compression of mortality were underway during that period, although the latter occurred at a slower pace since 1960. Ouellette and Bourbeau (2011) focus on changes in age at death distributions and mortality compression at older ages in the following low-mortality countries: Canada (1921–2007), the U.S. (1945–2007), France (1920–2009), and Japan (1947–2009). They find that similarly to Japanese and French females, old-age mortality compression no longer seems to be occurring among Canadian females for the most recent years. Instead, the female adult age at death distribution in Canada has been shifting to higher ages over time while maintaining an intact shape more recently, thus providing additional support to the *shifting mortality scenario*[3] first described by Kannisto (1996) and Bongaarts and Feeney (2002, 2003), and furthered by Bongaarts (2005). Canadian males were, however, still involved in reducing variability of deaths at advanced ages.

Study objectives

Given the absence of studies on regional variations in Canada with respect to the rectangularization of the survival curve and the compression of mortality, the present work aims mainly at shedding light on these topics at the level of Canadian provinces. In that sense, it can therefore be seen as a continuation of previous work by Nagnur (1986), Martel (2002), Martel and Bourbeau (2003), and Ouellette and Bourbeau (2011)—especially the latter, since our focus here is on adult and old-age mortality differentials. Furthermore, our approach also rests upon the flexible smoothing P-spline method; namely, we provide a generalization of the approach introduced by Ouellette and Bourbeau (2011) in order to make it particularly well-suited to the study of small populations.

Thus, the first objective of this paper is to highlight general converging and diverging adult mortality patterns among Canadian provinces using sex-specific smoothed mortality ratio surfaces for each province. This will allow us to gather high-level information on regional disparities among Canadian adults.

2. The late-life modal age at death, to be distinguished from the early-life modal age at death (at age 0), is found at older ages and corresponds to the age at which the largest number of deaths occurs. In other words, it is the most common age at death among adults.
3. The shifting mortality scenario suggests that adult mortality progressively shifts to higher ages over time, but that the mortality profile by age remains intact. Thus, under such a regime, the adult age at death distribution moves towards higher ages while keeping an intact shape.

Then, the second objective is to investigate more specifically the disparities among elderly populations of the various provinces and to compare their old-age mortality compression situations. Two-dimensional smoothed adult age at death distributions by sex will thus be used to monitor changes in the modal age at death and in the standard deviation of ages at death above the mode over time in each Canadian province. Province-specific time trends in the modal age at death will help us answer questions such as *What is the most common age at death in a given year in each province? How did it evolve over time? How does it differ from one province to another?* On the other hand, time trends in standard deviation above the mode will provide answers to complementary, yet equally important questions such as *What is the degree of variability in the age at death among elderly individuals in each province? How has it changed historically? How does it differ across Canadian provinces?*

Data

The data for this study were taken from the Canadian Human Mortality Database (CHMD 2010), which gathers detailed Canadian mortality and population data at the provincial and territorial level. This database is a unique source of information that allows us to study regional mortality patterns in Canada with great detail and accuracy. Indeed, as a satellite project of the well-known and widely used Human Mortality Database (2010) which includes 37 countries and areas, the CHMD also rests upon sophisticated demographic methods and high quality standards.[4]

Regarding data collection in Canada, there are currently thirteen systems of vital event registration in the country, that is, one for each province and territory. Most of them have been registering births and deaths since the 1920s, except for Newfoundland and Labrador, which joined the Canadian Confederation in 1949, and the territories, where registration coverage before 1950 tends to be incomplete. While the various provincial and territorial administrations are in charge of collecting the data, Statistics Canada is responsible for publishing them.

Although data included in the CHMD relies on thirteen different administrations, each with their own set of principles and practices that is subject to change over the years, there is no evidence of systematic discrepancies large enough to influence trends over time or between provinces and territories (Bourbeau et al. 2010). Indeed, registration problems are rather punctual, well known, and have been taken care of in the CHMD. Furthermore, missing information on deaths or information that needs to be corrected also arises, but concerns less than 1% of deaths each year (Bourbeau and Ouellette 2010).

Still, it should be mentioned that Kannisto (1994) and Kannisto et al. (1994) classified Canada as a country where data at older ages (i.e., 80 years and over) were of inferior quality compared to the Nordic countries, several European countries, and Japan. However, these analyses of Canadian data rested upon published data exclusively.[5] According to Bourbeau and Lebel (2000), such data were insufficient to draw the entire true picture in terms of data quality and mortality measurement at advanced ages in Canada. Bourbeau and Lebel's study uses more detailed data (unpublished data on deaths and population counts) and suggests that data quality in Canada, although not perfect, is quite good until age 100. The main data problems encountered in their analysis were among centenarians, more so for males than females, where age overstatement and errors in census age declarations were indeed more frequent. They also found that since 1951, the quality of data on deaths at ages 100 and older has generally improved over time. In fact, recent research on the validation of ages at death (using birth certificates from parish registers) among French-Canadian centenarians born in the province of Quebec who died between 1970 and 2007 demonstrates that even at the oldest ages, data quality is very good for this subgroup of the Canadian population (Beaudry-Godin 2010).

As a precautionary measure, given that results pertaining to the standard deviation of ages at death above the mode could be influenced by age overstatement problems among centenarians, findings for the first half of the 20th century, especially for males, should be interpreted with greater care. On the other hand, results for the modal age

4. To date, the CHMD is the only regional database to operate under the umbrella of the Human Mortality Database. The CHMD project is in fact a collaborative achievement of the Mortality and Longevity Research Team at the Département de Démographie, Université de Montréal, together with demographers at the Max Planck Institute for Demographic Research and the Department of Demography, University of California, Berkeley.
5. Access to Canadian data unpublished by Statistics Canada generally requires a special authorization from each province and territory. For example, deaths by single year of age over the entire age range for Canada are not published by Statistics Canada; neither are deaths by birth cohort.

at death, a robust measure in the presence of outliers, are likely to be insensitive to age overstatement issues among those aged 100 and over.

For the purpose of this paper, we extracted from the CHMD observed deaths counts and exposure data by sex, single year of age, and single calendar year for each Canadian province. Data below age ten were not extracted, given our aim to focus on adult and old-age mortality. Furthermore, data covers the 1930–2007 period,[6] except for the province of Newfoundland and Labrador, for which data starts in 1949. The Canadian territories, namely Yukon, Northwest Territories, and Nunavut, were excluded from our analysis mainly because of their very small population size. Indeed, there are so few deaths occurring each year for these territories[7] that even with a smoothing method well-suited for handling small populations, it would have been difficult to conduct an analysis by sex on these data. Furthermore, above the population size limitation, Canadian territories have systematically exhibited important mortality differentials from Canadian provinces over the years, mainly because of their distinct socioeconomic and cultural environment. Mortality and health disparities between Canadian territories and provinces are substantially lower today than in the 1950s, but worrying nonetheless (Prud'homme 2007; Veugelers et al. 2001). Canadian territories would thus deserve to be part of a study that specifically addresses their mortality developments over the second half of the 20th century.

Methods

Smoothing mortality data with P-splines

Recently, Ouellette and Bourbeau (2011) introduced the use of a one-dimensional smoothing approach, usually known as the P-spline method (Eilers and Marx 1996), to monitor with great precision changes in the age at death distribution over time in low-mortality countries at older ages. In the present paper, we rely on the two-dimensional version of this method in order to obtain sex-specific smoothed mortality surfaces and adult age at death distributions for each of the ten Canadian provinces. The P-spline method in two dimensions is particularly useful when dealing with small populations such as Canadian provinces, because it models mortality change over age and over time simultaneously. Therefore, the model consistently uses data information on neighboring ages and years, and punctual variations due to small numbers of deaths or exposures are less likely to distort the outcome (Camarda 2012). The details on the P-spline method in general, and in the context of mortality data specifically, are described in Appendix A.

In comparison to other methods for modeling mortality over age and over time, such as the Brass method (Brass 1971), the age-period-cohort model (Clayton and Schifflers 1987), and Lee-Carter approaches (Lee and Carter 1992; Brouhns et al. 2002), the P-spline method does not make any rigid assumption about the functional form of the mortality surface (Camarda 2008). This leads to a very fine expression of the underlying mortality patterns over age and time described by the data.

We used the *MortalitySmooth* package by Camarda (2009, 2012) for the statistical programming environment R (R Development Core Team 2010) to obtain smoothed two-dimensional forces of mortality by sex and province. Indeed, this package is precisely intended to perform P-spline smoothing in one and two dimensions for mortality data. The sex-specific smoothed death rates by age and time obtained for each province will be presented on shaded contour maps (Vaupel et al. 1997) in the Results section below. These graphical representations will also be specifically referred to as *mortality surfaces* in the remainder of the text. Mortality surfaces can summarize a considerable volume of data on a single graph and thus are very useful for looking at mortality trends. Given our aim to analyze mortality differences between Canadian provinces, sex- and province-specific *mortality ratio surfaces*, specifically adapted for this task, will also be presented in the Results section. These mortality ratio surfaces consist in shaded contour maps displaying ratios of sex-specific smoothed death rates by age and time for each province to sex-specific smoothed death rates by age and time for Canada, all provinces and territories combined. In the meantime, the next subsection explains how smoothed two-dimensional age at death distributions are computed and employed in this paper.

6. Given that some minor irregularities may have occurred in the beginnings of vital registration in Canada, data from the 1921–29 period were not included in our analysis.
7. In 2007, for example, Yukon, Northwest Territories, and Nunavut each recorded less than 200 deaths in total (Statistics Canada 2010: 20).

Computation and use of smoothed age at death distributions

The force of mortality, the survival function, and the density function are three particularly useful functions for describing mortality distributions. They also share the following interesting property: if one of these three functions is known, the remaining two can be uniquely determined. Thus, from sex-specific smoothed forces of mortality for each province, we can compute the corresponding smoothed density functions describing the age at death distributions in these populations.

Let $\mu(a,y)$ denote the force of mortality, a continuous function of age a and time y in a given population. Similarly, let $S(a,y)$ and $f(a,y)$ be the survival function and density function, respectively, which are also continuous functions of age and time. Given the usual correspondence between these three functions (Klein and Moeschberger 1997: 21–32), we have

$$\begin{aligned} f(a,y) &= \mu(a,y)\, S(a,y) \\ &= \mu(a,y) \exp\left(-\int_0^a \mu(u,y)\, du\right). \end{aligned} \tag{1}$$

Therefore, if the smoothed force of mortality $\hat{\mu}$ is known, then the corresponding smoothed density function $\hat{f}$ describing the two-dimensional age at death distribution can be computed using equation (1) and standard numerical integration techniques.

In order to monitor changes in the central tendency and old-age dispersion of adult deaths over time and across provinces, we used the following summary measures, respectively: the modal age at death and the standard deviation of individual life durations above the modal age at death. Inspired by Lexis's concept of normal life durations (Lexis 1877, 1878), Kannisto (2000, 2001) suggested this set of measures to focus specifically on changes in the age at death distribution occurring at older ages. This approach, with its emphasis on the modal age at death, has drawn much attention and was elaborated further in several recent studies (Cheung et al. 2005, 2008, 2009; Cheung and Robine 2007; Canudas-Romo 2008; Thatcher et al. 2010; Ouellette and Bourbeau 2011; Brown et al. 2012).

From the smoothed density function $\hat{f}$ describing the two-dimensional age at death distribution, we first computed the modal age at death using

$$\widehat{M}(y) = \max_a \hat{f}(a,y),$$

and then the standard deviation of individual life durations above the modal age at death (referred to hereafter as *standard deviation above the mode*) using

$$\widehat{SD(M+)}(y) = \sqrt{\frac{\int_{\widehat{M}(y)}^{\infty} \left(a - \widehat{M}(y)\right)^2 \hat{f}(a,y)\, da}{\int_{\widehat{M}(y)}^{\infty} \hat{f}(a,y)\, da}}.$$

Since $\hat{f}$ is a function of ages and years, $\widehat{M}$ and $\widehat{SD(M+)}$ are functions of years.

For a more exhaustive monitoring of changes in variability in old-age mortality, $\widehat{SD(M+)}$ could be supplemented by a set of mode-based quantile functions (Ouellette 2011). Indeed, mode-based quantiles provide greater flexibility than $\widehat{SD(M+)}$; they allow us to focus on changes in variability which took place in any given neighborhood above $\widehat{M}$, no matter how narrow or wide. However, in a first attempt to highlight regional disparities in old-age mortality variabity among Canadian provinces, $\widehat{SD(M+)}$ is the most suitable measure.

Given that provinces in Canada are very distinct in population size, 99% confidence intervals for $\widehat{M}(y)$ and for $\widehat{SD(M+)}$ (y) were also computed for each year $y \in \{1930, 1931, \ldots, 2007\}$ to facilitate comparisons across provinces. The details of the residual bootstrap method addressed by Koissi et al. (2006) and used in the present study to construct these confidence intervals are described in Appendix B.

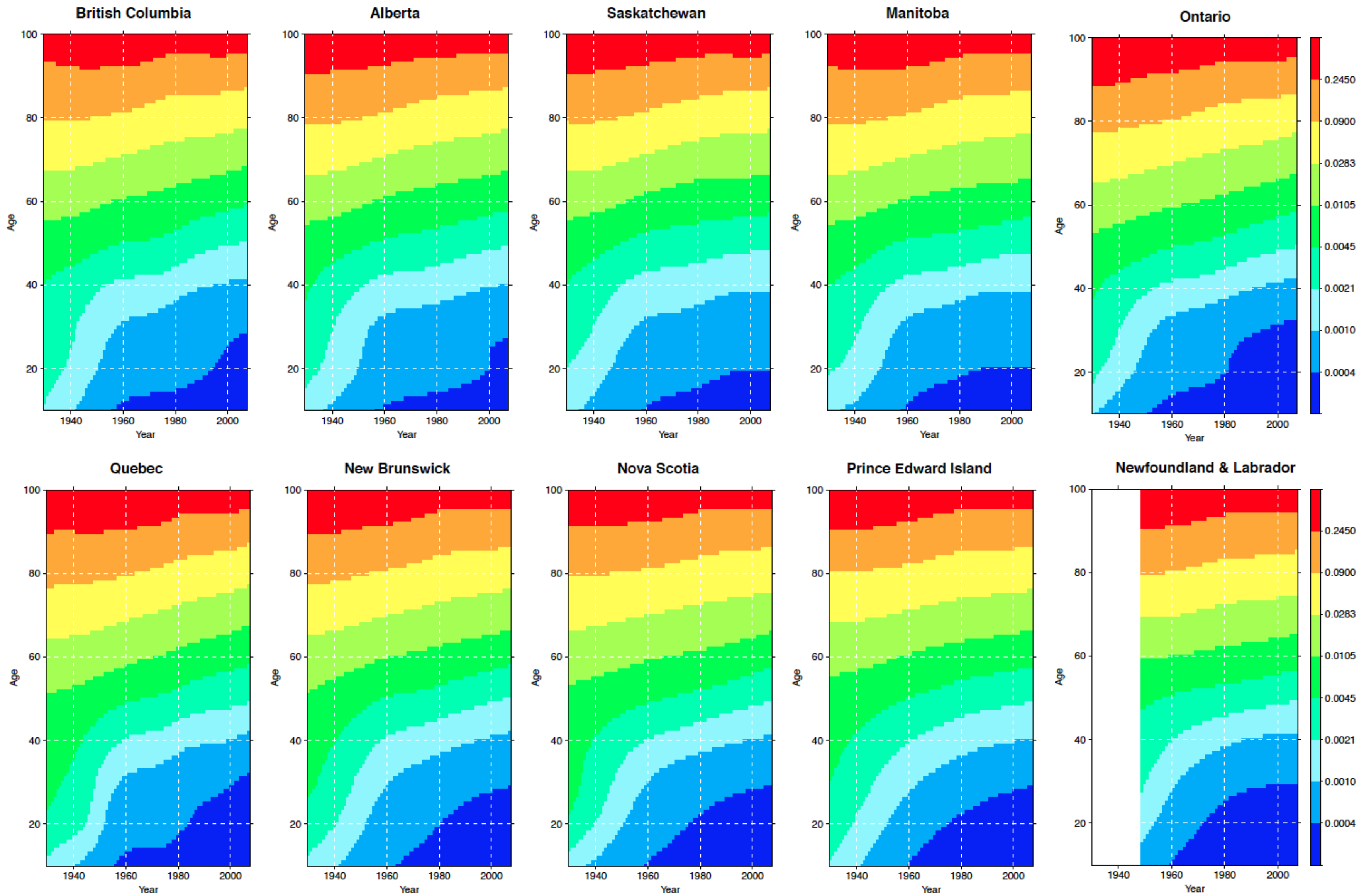

Figure 1. Province-specific smoothed mortality surfaces based on fitted death rates with two-dimensional Poisson P-splines, Females, Canada, 1930 to 2007.

Note: Extra-Poisson variation (overdispersion) was detected in Ontario and accounted for in the smoothing procedure.

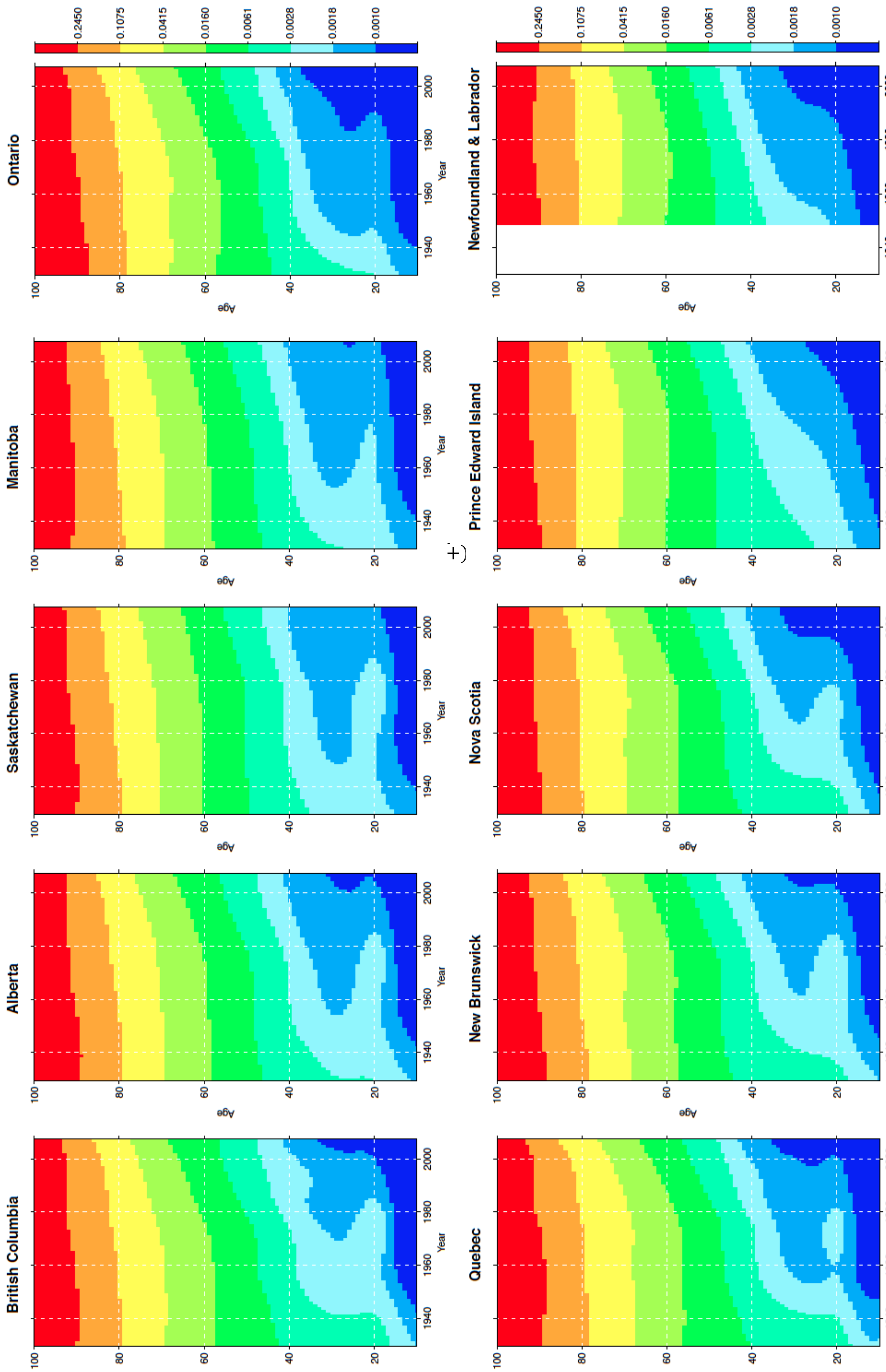

Figure 2: Province-specific smoothed mortality surfaces based on fitted death rates with two-dimensional Poisson P-splines, Males, Canada, 1930 to 2007.

Note: Extra-Poisson variation (overdispersion) was detected in Ontario and accounted for in the smoothing procedure.

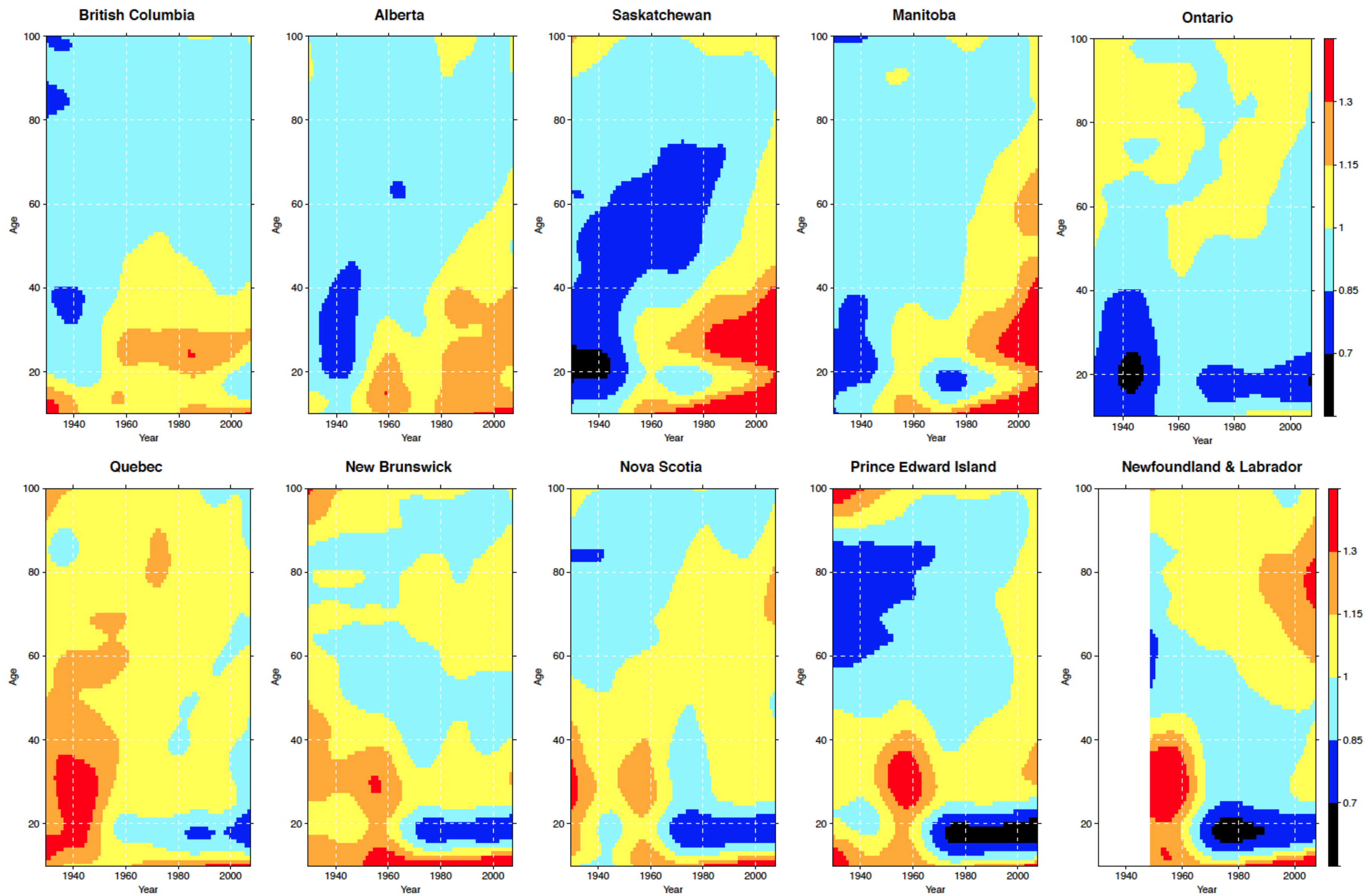

Figure 3: Province-specific smoothed mortality ratio surfaces (with respect to total female Canadian population), based on fitted death rates with two-dimensional Poisson P-splines, Females, Canada, 1930 to 2007.

Note: Extra-Poisson variation (overdispersion) was detected in Ontario and accounted for in the smoothing procedure.

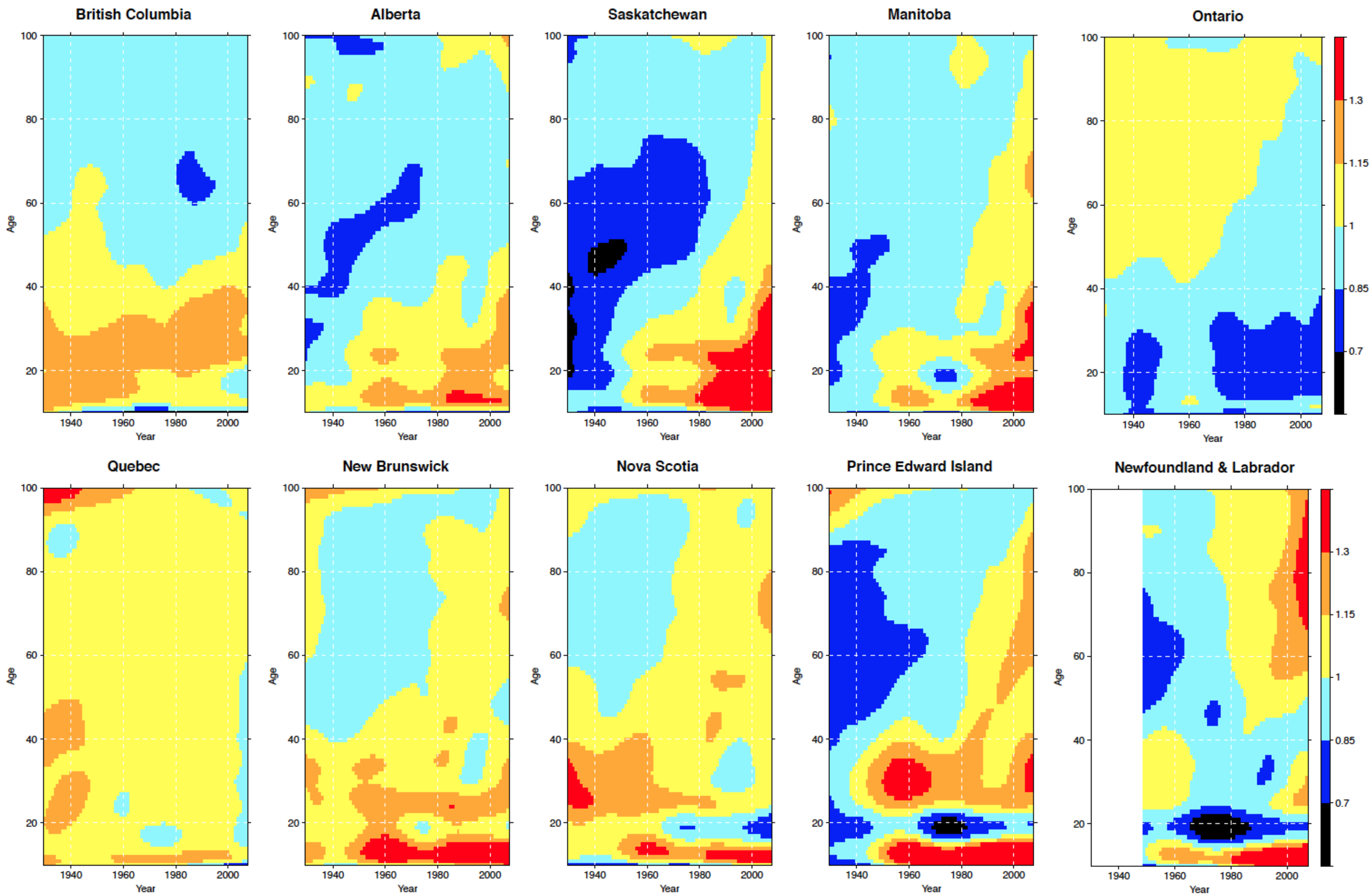

Figure 4: Province-specific smoothed mortality ratio surfaces (with respect to total male Canadian population) based on fitted death rates with two-dimensional Poisson P-splines, Males, Canada, 1930 to 2007.

Note: Extra-Poisson variation (overdispersion) was detected in Ontario and accounted for in the smoothing procedure.

Results

Uncovering provincial variations via smoothed mortality ratio surfaces

Figures 1 and 2 show province-specific smoothed mortality surfaces for females and males, respectively. To our knowledge, these are the very first smoothed mortality surfaces for Canadian provinces to be published. Each of these shaded contour maps displays smoothed death rates over age and time, and thus summarizes a great amount of information on a single graphic representation. The scales on the right-hand side of Figures 1 and 2 show how smoothed death rates were first partitioned into nine groups for each sex. Death rates belonging to the same group were then assigned the same color. For example, female death rates below 0.0004 were all plotted in dark blue on the female mortality surfaces and they represent the lowest level of mortality on the surface. Among males, this group includes all death rates below 0.001. Likewise, female death rates higher than 0.245 correspond to the highest level of mortality on the surface and were all plotted in red. Among males, this group also starts at 0.245.

From Figures 1 and 2, mortality improvements between 1930 and 2007 at most adult ages and across all provinces for both sexes are evident. Indeed, the upward trend followed by most color bands indicates that low levels of mortality are progressively associated to a wider range of ages as the years go by. We also note that reductions in mortality at higher ages, say above 55 years old, roughly began in the 1970s among males, while they were already apparent in the 1930s among females. Furthermore, in geographical terms, provinces that belong to the same great-Canadian region—that is, the Western provinces (British Columbia, Alberta, Saskatchewan, Manitoba), Central provinces (Ontario, Quebec), and Atlantic provinces (New Brunswick, Nova Scotia, Prince Edward Island, Newfoundland and Labrador)—tend to reveal similar mortality patterns. Nevertheless, substantial variations exist within these broad regions, as we will discuss shortly.

Although the smoothed mortality surfaces presented in Figures 1 and 2 are very helpful for summarizing mortality trends over age and time, they are not as suitable for analyzing mortality differences between Canadian provinces. In fact, the smoothed mortality ratio surfaces presented in Figures 3 and 4 are much more appropriate for such a task. In each of these sex- and province-specific mortality ratio surfaces, every point represents the ratio of the original smoothed mortality surface to the smoothed sex-specific mortality surface for Canada, all provinces and territories combined. Using a common denominator for the various provinces guarantees that mortality ratio surfaces remain comparable for a given sex. Scales on the right-hand side of Figures 3 and 4 show how ratios of smoothed death rates were partitioned for display on the surfaces. Shades of light blue, dark blue, and black, respectively, indicate progressively lower levels of mortality in a province compared to the corresponding (female or male) total Canadian population. Similarly, shades of yellow, orange, and red, respectively, reveal increasingly higher mortality disadvantage for a province with respect to the Canadian population.

Starting with the Western provinces, the situation in Saskatchewan and Manitoba differs quite substantially from the one in British Columbia, while Alberta occupies an intermediate position. Between 1930 and 1960, mortality at almost every adult age below 70 in Saskatchewan was the lowest of all Western provinces for each sex. In fact, Saskatchewan even maintained lower mortality for adults aged between 40 and 70 up to the 1980s. Nonetheless, this advantage below age 70 was then progressively replaced by rather disadvantageous mortality conditions which lasted until the end of the study period. Below age 40, in particular, mortality improvements in Saskatchewan have likely been relatively slow since the 1960s, given the growing predominance of yellow, orange, and ultimately red shaded areas on both female and male surfaces. The overall pictures for each sex in Manitoba are highly similar to those of its immediate western neighbor, although slightly more moderate. Indeed, most features discussed for Saskatchewan can be found on the mortality ratio surfaces for Manitoba, but shades of dark blue or black and orange or red are all confined to narrower areas.

In British Columbia, female mortality below age 30 during the 1930s and 1940s was generally higher than in the other Western provinces. However, mortality reductions at these ages occurred at a rapid pace afterwards, and by the mid-1980s and onwards, British Columbia females were the most advantaged ones within the Western provinces. Such an initial mortality disadvantage in the 1930s with respect to other Western provinces is also visible among British Columbia males below age 70. Although it lasted longer than among females, it almost completely disappeared over time. Another aspect emerging from British Columbia mortality ratio surfaces for each sex is that, unlike in any other Canadian province (non-Western provinces especially), moderately low levels of mortality were consistently recorded at the oldest ages, say 75 and above, between 1930 and 2007.

Alberta occupies an intermediate position among Western provinces; depending on the perspective (which ages and years considered), female and male mortality ratio surfaces are either closer to the ones for British Columbia or for Saskatchewan and Manitoba.

Within the Central provinces, the level of mortality in the 1930s, 1940s, and 1950s was by far higher in Quebec than in Ontario at almost every adult age below 70 among females. Similarly, males in Quebec experienced greater mortality than males in Ontario below age 50, but the discrepancies between the two provinces were less striking than among females during those years. Throughout most of the subsequent years, females of all ages above 10 in Ontario generally maintained an advantage over their Quebec counterparts. However, in the latest decade, the gap between the two provinces was substantially narrower. Indeed, based on the mortality surfaces from Figure 1, thanks to an intensive phase of reductions in mortality at ages below 50 during the first half of the 20th century, plus a sustained high rate of progress at ages above 50 throughout the study period, Quebec females have now almost caught up with those of Ontario. Strong reductions in mortality also occurred among males in Quebec (see Fig. 2) and led them closer to their male counterparts in Ontario in the last years studied (see Fig. 4).

Among the Atlantic provinces, the smoothed mortality ratio surfaces from Figures 3 and 4 for New Brunswick and Nova Scotia share many similarities overall for each sex. Still, New Brunswick shows a slightly higher female mortality at several ages above 10 compared to Nova Scotia between 1930 and the 1960s, especially above age 65. Similarly, males in New Brunswick experienced higher mortality above age 60 in the first decade of the period under study, but managed to record lower mortality levels than Nova Scotia at these ages during the 1970s. With respect to mortality around age 20, males in Nova Scotia have had, on the other hand, an advantage over their male counterparts from New Brunswick for more than four decades now.

In Prince Edward Island, females and males seemed to enjoy the lowest mortality of all Atlantic provinces at most ages below 90 during the 1930s. Moreover, this advantage even extended to the 1960s between ages 50 and 85. However, given the small size of Prince Edward Island's population, these results are based on scarce information and should be interpreted with greater care. In fact, as the years went by afterwards, shades of yellow, orange, and red progressively covered a broader range of ages, especially among males. In the last decade, male mortality in Prince Edward Island between ages 50 and 90 was even higher than in New Brunswick and Nova Scotia, suggesting that reductions in male mortality in this province have been much slower than in New Brunswick and Nova Scotia since the 1950s (see Fig. 2). Newfoundland and Labrador also shows a gradual yet sustained deterioration of its advantageous mortality conditions at ages above 50 throughout the second half of the 20th century for each sex. Furthermore, in the latest years studied, Newfoundland and Labrador was clearly the most disadvantaged Canadian province in terms of old-age mortality.

Thus, adult and old-age mortality differentials do exist among Canadian provinces and have been roughly revealed by the smoothed mortality ratio surfaces described above. The next subsection delves one step further into these differentials and examines changes in smoothed province-specific adult age at death distributions over time. Given the summary measures favoured to monitor these changes—namely, the modal age at death and the standard deviation above the mode—our focus is on the elderly population specifically rather than on all individuals aged 10 and above.

Summarizing changes in smoothed age at death distributions over time across provinces

Figure 5 presents estimated time trends in the modal age at death by province for each sex, computed from their respective two-dimensional age at death distribution. Among females, all provinces showed an upward trend throughout most of the period under study. These upward trends were in fact almost linear between 1940 and 1980, and in the last two decades. In the early 1930s, female modal age at death estimates for most provinces lay between 78 and 80 years.[8] At the turn of the 21st century, the female modal age at death values had increased by about 10 years in most provinces, and ranged from 87 to 90 years. Among Western provinces, females displayed very similar results from 1930 to 2007. Indeed, 99% confidence intervals for their modal age at death estimates reveal that significant differences mainly occurred only between Saskatchewan and Manitoba during the 1990s (see Table B.1 in Appendix B for 99% confidence intervals results). Throughout the period under study, females in Western provinces enjoyed high modal age at death outcomes compared to those of other Canadian provinces. In fact, from the 1960s until very recently, the confidence interval yield for female modal age at death in British Columbia, Alberta, and Saskatchewan was significantly higher than in both Central provinces and most Atlantic provinces. Since 2000, however, modal age at death differences between Western provinces and Central provinces have been no longer statistically different. This

8. The female modal age at death estimates for Prince Edward Island between 1930 and 1960 are higher than in any other Canadian province. However, due to the small population size of Prince Edward Island, its 99% confidence intervals for modal age at death estimates are wide (see Table B.1 in Appendix B). Thus, despite appearances, the female modal age death for Prince Edward Island in the 1930s is not significantly higher than in British Columbia.

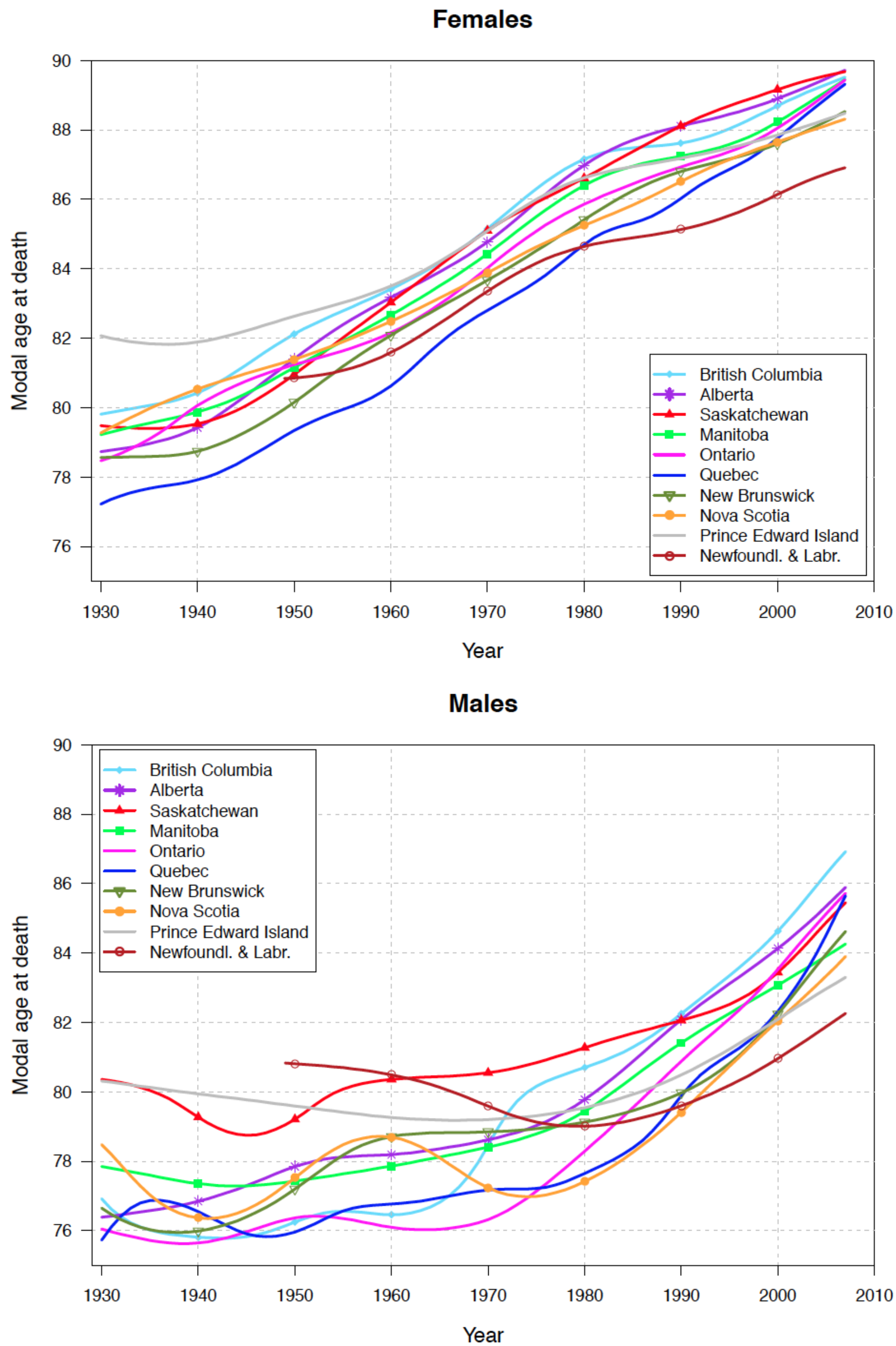

Figure 5. Modal age at death by province and sex estimated from two-dimensional smoothed age-at-death distribution, Canada, 1930–2007.

is particularly interesting for Quebec females, who consistently recorded significantly lower modal age at death results than most of their female counterparts between 1930 and 1990. In contrast, the situation among females in Atlantic provinces has deteriorated over time. Moreover, in the last few years under study, the female modal age at death in these provinces was significantly lower than in any other Western or Central province.

Among males, trends in modal age at death over time presented in the bottom panel of Figure 5 reveal that sustained increases did not start much before the late 1960s or even early 1970s in Canadian provinces. Indeed, from 1930 up to the 1960s, the modal age at death trend for males either stagnated or rippled in the high-70s range, probably because mortality reductions at ages above the mode, essential for its increase (Canudas-Romo 2010), were limited during that period of time. Afterwards however, substantial increases in the male modal age at death were recorded in all provinces, and 2007 estimates lay between 82 and 87 years. The case of British Columbia males stands out because in the early 1960s, their modal age at death estimates were still among the lowest, but since 1990 they have consistently been the highest. According to the 99% confidence intervals for male modal age at death estimates in Table B.1, British Columbia results have been significantly higher than in each Atlantic province since about 1980, and in Central provinces between 1990 and 2005. Within the Western provinces, significant differences are scarce. Indeed, they mainly occurred between Saskatchewan and Manitoba during the 1960s and 1970s, and between British Columbia and Manitoba in the latest years. Similarly to British Columbia, the Central provinces were also displaying some of the lowest modal age at death estimates for males between 1930 and the 1960s, and made impressive progress afterwards. Among the Atlantic provinces, despite increases in the male modal age at death since the 1970s, Newfoundland and Labrador, Prince Edward Island, and Nova Scotia have been showing the lowest results of all Canadian provinces since 2000. However, differences with New Brunswick, Quebec, Manitoba, and Saskatchewan were not systematically significant.

Another feature of Figure 5 worth mentioning concerns the sex differences in modal age at death. Throughout the period studied, females in each province enjoyed almost systematically higher results than their male counterparts. Similarly to the sex gap in life expectancy at birth, the gap in modal age at death for most provinces increased steadily between 1930 and the 1970s, but then began to contract and kept narrowing thereafter. In 2007, sex differentials in modal age at death across Canada ranged from 2.6 years in British Columbia to about 5.2 years in Prince Edward Island and Manitoba. Furthermore, according to the confidence intervals provided in Table B.1, 2007 modal age at death results for females and males in all provinces were still statistically different.

Figure 6 displays sex- and province-specific time trends in the standard deviation above the modal age at death computed from two-dimensional age at death distributions. Among females, each province recorded lower results at the end of the period covered by this study than at the beginning. In other words, old-age compression of mortality occurred between 1930 and 2007 among these females. However, the trends did not decline steadily over the entire period. Indeed, the 1960s and 1970s were years of slow decline for most provinces. Furthermore, since 2000, the female standard deviation above the mode in Western provinces (except Manitoba) and Central provinces has decreased at a much slower pace than in previous years, stopped decreasing, or even increased slightly. Evidence from the Atlantic provinces instead shows that old-age compression of mortality continues among females, because their standard deviation above the mode did not stop declining in recent years. This is perhaps related to the fact that the level of female old-age mortality compression has remained higher in most Atlantic provinces than in other Canadian provinces since 2000. Indeed, the 99% confidence intervals for the female standard deviation above the mode estimates (see Table B.2) reveal that while significant differences across provinces were rather scarce between 1930 and 1970, they became more common afterwards, especially between Atlantic and Western provinces.

Among males, we notice from the bottom panel of Figure 6 that the province-specific standard deviation above the mode estimates were not necessarily lower in 2007 than in 1930. In other words, unlike among Canadian females, old-age compression of mortality did not occur in all ten provinces over the period as a whole for males. However, if we limit ourselves to the period starting with the onset of the modal age at death increases for Canadian males—that is, around 1970—then it generally coincides with episodes of old-age mortality compression in all provinces. Indeed, male standard deviation above the mode estimates decreased quite steadily over time after 1970. Even in recent years, the pace of decline did not slow down, except perhaps in Ontario. Since 1990, male standard deviation above the mode estimates in Atlantic provinces have been among the highest in Canada. In contrast, Western provinces generally enjoyed the lowest estimates for males, especially British Columbia and Alberta. Indeed, for several years in the 1990s and early 2000s, male standard deviation above the mode estimates in Prince Edward Island, and Nova Scotia were significantly higher than in British Columbia and Alberta (see Table B.2). Differences between New Brunswick and British Columbia were also significant during most of the 1990s.

In terms of sex differences in standard deviation above the mode, Figure 6 reveals that females in each province usually recorded lower results than males. Up to the 1970s, the sex gap in standard deviation above the mode tended to increase in most provinces, but it has been narrowing ever since. In 2007, results for females and males in almost all provinces were no longer statistically different (see Table B.2).

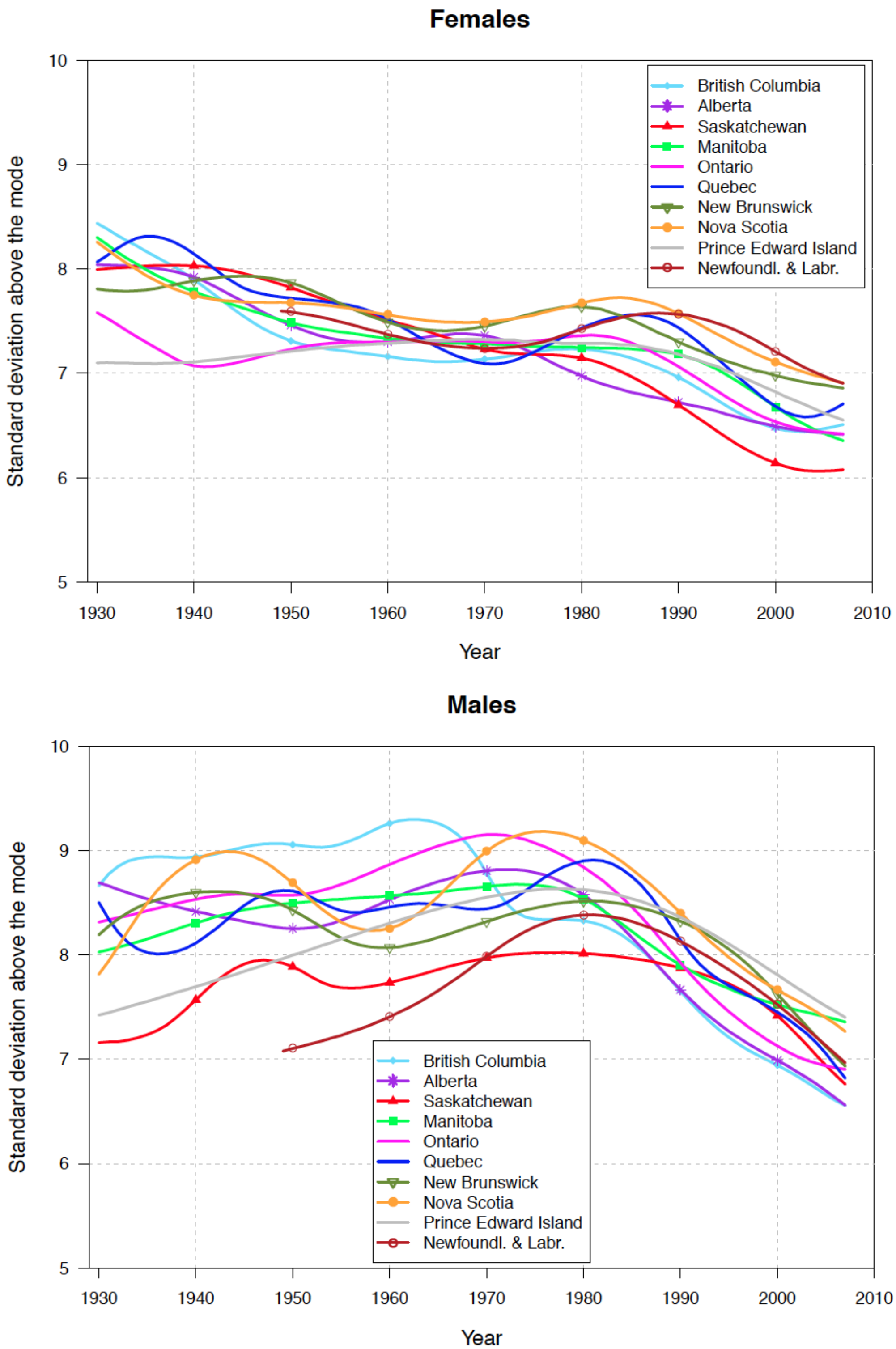

Figure 6. Standard deviation above the modal age at death by province and sex estimated from two-dimensional smoothed age-at-death distribution, Canada, 1930–2007.

Discussion

This paper explored differences in adult mortality among Canadian provinces between 1930 and 2007, and focused notably on old-age mortality compression, a topic that has not been investigated yet at the regional level in this country. Our analysis revealed that provincial disparities in adult mortality in general, and among the elderly

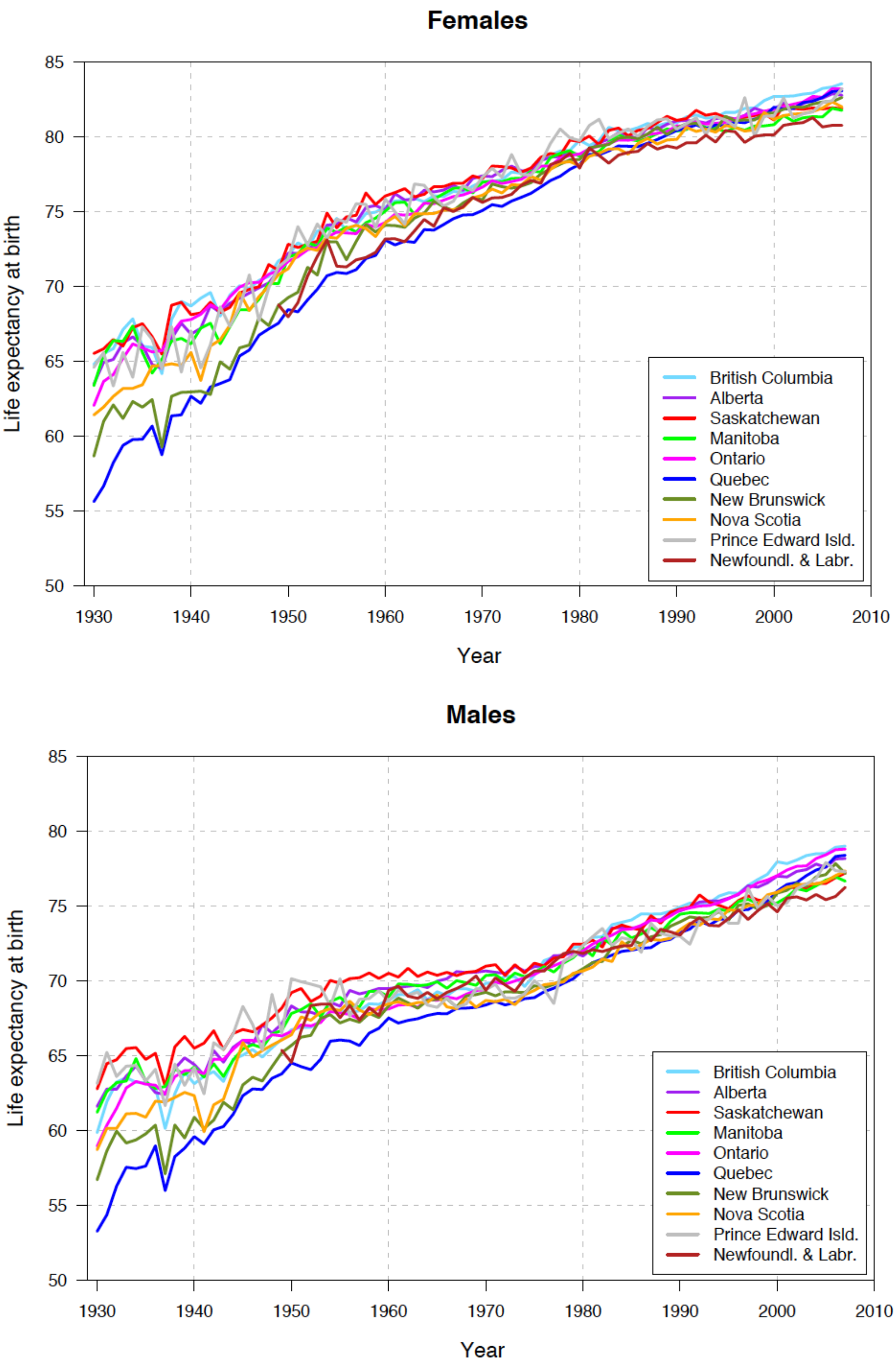

Figure 7: Life expectancy at birth by province and sex, Canada, 1930 to 2007.

population in particular, are substantial in Canada. Moreover, these disparities have barely reduced over time, despite the great mortality improvements recorded within each of these provinces since the beginning of the 20th century.

Province-specific smoothed mortality ratio surfaces by sex (see Fig. 3 and 4) illustrated that throughout the period under study, females and males in Saskatchewan and Manitoba progressively lost their mortality advantage over those of other Canadian provinces at ages below 75. In recent years, mortality decline among young adults (below age 40) has been particularly slow in these two provinces. At older ages, Newfoundland and Labrador, and to a lesser extent Prince Edward Island, displayed the worst mortality conditions for each sex in the latest years. In contrast, females and males in Quebec have been the most successful at lowering mortality at almost every age above 10, especially since the 1950s. Finally, British Columbia is the only Canadian province that consistently maintained moderately low levels of mortality between 1930 and 2007 at the oldest ages (75 and above) for each sex.

These findings are consistent with those of previous studies on mortality disparities in Canada (Manuel and Hockin 2000; Prud'homme 2007) where the relatively weak recent mortality improvements in Saskatchewan were underlined and the rise of Quebec in the provincial ranking emphasized. These studies also identified British Columbia as the new leader and reference in terms of mortality in Canada. Furthermore, these recent trends in regional mortality differentials among adults in Canada are in line with findings emerging from analyses of the prevalence of risk behaviors and health conditions for cardiovascular diseases (PHAC 2009). Indeed, the lowest 2007 prevalence figures for smoking, physical inactivity, and obesity among Canadian provinces systematically belong to British Columbia. Alberta has the lowest prevalence of high blood pressure, but British Columbia follows closely. On the opposite side of the country, Atlantic provinces consistently display higher prevalence results than the Canadian average. Moreover, prevalence of obesity and high blood pressure are both particularly high in Newfoundland and Labrador.

Time-trends results in modal age at death by province and sex (see Fig. 5) further helped us appreciate the magnitude of the disparities across the provincial elderly populations. These modal age at death trends are in fact a great addition to the usual life expectancy at birth trends (see Fig. 7) often reported to discuss regional mortality differentials in Canada. Disparities in life expectancy at birth across Canadian provinces have been largely reduced over the 20th century, because all provinces made remarkable progress in terms of infant mortality and infectious as well as parasitic diseases during that period. However, Figure 7 hides the fact that old-age mortality disparities have barely been reduced among Canadian provinces between 1930 and 2007. This part of the story, rather, emerges from Figure 5 because unlike life expectancy at birth, which is very sensitive to improvements in mortality among infants and children, modal age at death is sensitive to mortality changes occurring among the elderly population. The range of values taken over time for province-specific modal ages at death is thus much narrower than for life expectancies at birth, and therefore, regional changes in old-age mortality can be specifically addressed. Indeed, Figure 5 shows that in the last decade of our study, as the modal age at death was increasing very rapidly in Quebec and Ontario, these Central provinces were actually catching up with Western provinces, while the lag of Atlantic provinces was becoming more and more pronounced. The situation in Newfoundland and Labrador is particularly worrying and deserves to be carefully monitored in upcoming years. Thus, the long-standing geographical mortality disparities favouring provinces in the Western part of Canada compared to those in the East hold among the elderly population. However, this might be subject to change, as Central provinces could perhaps take the lead in Canada in the near future.

The level of variability of age at death among older individuals is another great source of provincial mortality disparity in Canada (see Fig. 6). Again, reductions in these provincial differentials over time were very scarce between 1930 and 2007. Indeed, at the end of the period under study, Canadian provinces still showed a broad palette of levels of old-age mortality dispersion. The highest levels of standard deviation above the mode tended to be recorded in provinces belonging to the Eastern part of Canada, while provinces with the lowest levels belonged to the Western part of Canada. Males in Manitoba are perhaps the most obvious exception, as their level of old-age mortality dispersion remains high compared to the other Western provinces.

Cause decomposition methods for changes in modal age at death and standard deviation above the mode are likely to be helpful for understanding why provincial disparities in old-age mortality persist in Canada. Fortunately, research on the development of cause decomposition methods for the modal age at death has recently been initiated by Canudas-Romo (2011), and this should stimulate further research on such aspects of mortality analysis in the near future.

Regarding sex differences in mortality at older ages, trends over time for the sex gap in modal age at death for most Canadian provinces share many similarities with the better documented sex gap in life expectancy at birth. However, the explanations for narrowing of the sex gap in modal age at death are likely to differ from those for life expectancy at birth, because these two summary measures are of a distinctive nature. Decomposition analysis by cause of death for sex differentials in modal age at death in two points in time (e.g., 1975 and 2005) would shed additional light on this topic.

When Kannisto (2000, 2001) introduced the modal age at death and the standard deviation above the mode to summarize and monitor changes in age at death distribution at older ages, the rationale behind his set of indicators was very clear. Episodes of modal age at death increases, paralleled with decreases in the standard deviation above the mode, suggest that the distribution of ages at death at old ages is not solely moving towards higher ages. As Kannisto stated, "Instead, its right-hand slope is being flattened vertically, as if it was meeting an invisible wall" (Kannisto 2001: 169). On the other hand, when the mode increases and the standard deviation above the mode no longer declines, the distribution of ages at death at old ages is simply sliding towards higher ages, without changing its shape. Thus, each of these scenarios, respectively known as the *old-age compression of mortality scenario* and the *shifting mortality scenario*, may provide valuable indications on human longevity and possible limits to the human lifespan.

In the present study, evidence of the shifting mortality regime was recorded among females in most Western and Central provinces, but males in all provinces were still engaged in an old-age mortality compression regime. These results concur with those of Ouellette and Bourbeau (2011), where Canadian females in all provinces and territories were found to be heading towards a shifting mortality regime in upcoming years, while males did not show such evidence. These findings, showing evidence of the shifting mortality regime lately, call for further research on the health situation of the elderly populations in Canadian provinces. Under such a regime, it is assumed that adult mortality shifts to higher ages (Bongaarts 2005). However, whether morbidity and disability episodes also shift to higher ages, or rather occupy a larger proportion of the typical lifespan, remains unclear. Research on Canadian provincial disparities in terms of modal age at death, and dispersion above it by cause of death, should also be helpful in that sense.

Acknowledgements

This project was funded by the Social Sciences and Humanities Research Council of Canada. The authors are grateful to the Council for its support. The authors would also like to thank the three anonymous reviewers for their valuable comments and advices for revision.

References

Adams, O. 1990. Life expectancy in Canada: An overview. *Health Reports* 2(4):361–376.

Beaudry-Godin, M. 2010. *La démographie des centenaires québécois : validation des âges au décès, mesure de la mortalité et composante familiale de la longévité*. PhD diss. Université de Montréal, Département de démographie.

Bongaarts, J. 2005. Long range trends in adult mortality: Models and projection methods. *Demography* 42(1):23–49.

Bongaarts, J., and G. Feeney. 2002. How long do we live? *Population and Development Review* 28(1):13–29.

————. 2003. Estimating mean lifetime. *Proceedings of the National Academy of Science* 100(23):13127–13133.

Bourbeau, R. 2002. L'effet de la "sélection d'immigrants en bonne santé" sur la mortalité canadienne aux grands ages. *Cahiers québécois de démographie* 31(2):249–274.

Bourbeau, R., and A. Lebel. 2000. Mortality statistics for the oldest-old: An evaluation of Canadian data. *Demographic Research* 2(2).

Bourbeau, R., J. Légaré, and V. Émond. 1997. New birth cohort life tables for Canada and Quebec, 1801–1991 (Cat. No. 91F0015MPE1997003). Ottawa: Statistics Canada.

Bourbeau, R., and N. Ouellette. 2010. The Canadian Human Mortality Database: A project worth the challenges! Paper presented at the Third Human Mortality Database Symposium, Paris (June 17–19).

Bourbeau, R., N. Ouellette, and C. Lecours. 2010. La base de données sur la longévité canadienne : corrections apportées aux données sur les décès de 1921 à 2007. Université de Montréal, Département de démographie.

Brass, W. 1971. On the scale of mortality, in *Biological Aspects of Demography*, edited by W. Brass. London: Taylor & Francis, pp. 69–110.

Brouhns, N., M. Denuit, and J.K. Vermunt. 2002. A Poisson log-bilinear regression approach to the construction of projected lifetables. *Insurance: Mathematics & Economics* 31:373–393.

Brown, D.C., M.D. Hayward, J.K. Montez, R.A. Hummer, C. Chiu, and M.M. Hidajat. 2012. The significance of education for mortality compression in the United States. *Demography* 49(3):819–840.

Camarda, C.G. 2008. Smoothing Methods for the Analysis of Mortality Development. PhD diss. Universidad Carlos III de Madrid (Spain) and Max Planck Institute for Demographic Research (Germany).

————. 2009. MortalitySmooth: Smoothing Poisson counts with P-splines. R package version 1.0.

————. 2012. MortalitySmooth: A R package for smoothing Poisson counts with P-splines. *Journal of Statistical Software* 50(1).

Canudas-Romo, V. 2008. The modal age at death and the shifting mortality hypothesis. *Demographic Research* 19:1179–1204.

————. 2010. Time trends and record values for longevity. *Demography* 47(2):299–312.

————. 2011. Cause-decomposition of the change over time in the modal age at death. Paper presented at the annual meeting of the Population Association of America. Washington, DC (March 31–April 2).

Chen, J., R. Wilkins, and E. Ng. 1996. Health expectancy by immigrant status, 1986 and 1991. *Health Reports* 8(3):29–38.

Cheung, S.L.K., and J-M. Robine. 2007. Increase in common longevity and the compression of mortality: The case of Japan. *Population Studies* 61(1):85–97.

Cheung, S.L.K., J-M. Robine, E.J-C. Tu, and G. Caselli. 2005. Three dimensions of the survival curve: Horizontalization, verticalization, and longevity extension. *Demography* 42(2):243–258.

Cheung, S.L.K., J-M. Robine, and G. Caselli. 2008. The use of cohort and period data to explore changes in adult longevity in low mortality countries. *Genus* LXIV(1–2):101–129.

Cheung, S.L.K., J-M. Robine, F. Paccaud, and A. Marazzi. 2009. Dissecting the compression of mortality in Switzerland, 1876–2005. *Demographic Research* 21:569–598.

CHMD (Canadian Human Mortality Database). 2010. Université de Montréal (Canada), University of California, Berkley (USA), and Max Planck Institute for Demographic Research (Germany). Accessed via http://www.bdlc.umontreal.ca (data downloaded on [16/07/2010]).

Clayton, D., and E. Schifflers. 1987. Models for temporal variation in cancer rates II: Age-period-cohort models. *Statistics in Medicine* 6:469–481.

Currie, I.D., M. Durban, and P.H.C. Eilers. 2004. Smoothing and forecasting mortality rates. *Statistical Modelling* 4:279–298.

Eakin, T., and M. Witten. 1995. How square is the survival curve of a given species? *Experimental Gerontology* 30(1):33–64.

Efron, B., and R.J. Tibshirani. 1993. *An Introduction to the Bootstrap*. Monographs on Statistics and Applied Probability. London: Chapman & Hall.

Eilers, P.H.C., and B.D. Marx. 1996. Flexible smoothing with B-splines and penalties (with discussion). *Statistical Science* 11(2):89–102.

Field, N.C. 1980. Temporal trends and spatial patterns of mortality in Canada, in *Canadian Studies in Medical Geography*, edited by F. A. Barrett. Toronto: Department of Geography, Atkinson College, York University, pp. 32–58.

Fries, J.F. 1980. Aging, natural death, and the compression of morbidity. *New England Journal of Medicine* 303:130–35.

Goyder, J. 1993. The Canadian syndrome of polarities: An obituary. *Canadian Review of Sociology and Anthropology* 30(1):1–12.

Health Canada. 1995. Canadians and Heart Health, Reducing the Risk (Cat. No. H39-328/1995E). Ottawa: Minister of Supply and Services Canada.

Hill, G. 1993. The entropy of the survival curve: An alternative measure. *Canadian Studies in Population* 20(1):43–57.

Horiuchi, S. 2003. Interspecies differences in the life span distribution: Humans versus invertebrates, edited by J.R. Carey and S. Tuljapurkar. *Population and Development Review* 29(Supplement):127–151.

Human Mortality Database. 2010. University of California, Berkley (USA), and Max Planck Institute for Demographic Research (Germany). Accessible via http://www.mortality.org.

Jeune, B., and J.W. Vaupel 1995. *Exceptional Longevity: From Prehistory to the Present*. Odense: Odense University Press.

Kannisto, V. 1994. *Development of Oldest-Old Mortality, 1950–1990: Evidence from 28 Developed Countries*. Monographs on Population Aging 1. Odense: Odense University Press.

————. 1996. *The Advancing Frontier of Survival Life Tables for Old Age*. Monographs on Population Aging 3. Odense: Odense University Press.

————. 2000. Measuring the compression of mortality. *Demographic Research* 3(6).

————. 2001. Mode and dispersion of the length of life. *Population: An English Selection* 13(1):159–171.

————. 2007. Central and dispersion indicators of individual life duration: New methods, in *Human Longevity, Individual Life Duration, and the Growth of the Oldest-old Population*, edited by J-M. Robine, E. Crimmins, S. Horiuchi, and Y. Zeng. Netherlands: Springer, pp. 111–129.

Kannisto, V., J. Lauristen, A.R. Thatcher, and J.W. Vaupel. 1994. Reductions in mortality at advanced ages: Several decades of evidence from 27 countries. *Population and Development Review* 20(4):793–809.

Klein, J.P., and M.L. Moeschberger. 1997. *Survival Analysis: Techniques for Censored and Truncated Data*. New York: Springer.

Koissi, M-C., A.F. Shapiro, and G. Högnäs. 2006. Evaluating and extending the Lee-Carter model for mortality forecasting: Bootstrap confidence interval. *Insurance: Mathematics and Economics* 38:1–20.

Lee, R.C., and L.R. Carter. 1992. Modeling and forecasting U.S. mortality. *Journal of the American Statistical Association* 87:659–675.

Lexis, W. 1877. *Zur Theorie der Massenerscheinungen in der menschlichen Gesellschaft* [On the Theory of Mass Phenomena in Human Society]. Freiburg i.B.: Fr. Wagner'sche Buchhandlung.

————. 1878. Sur la durée normale de la vie humaine et sur la théorie de la stabilité des rapports statistiques. *Annales de démographie internationale* 2:447–460.

Lopez, A.D., and L.T. Ruzicka. 1983. *Sex Differentials in Mortality: Trends, Determinants and Consequences*. Canberra: Australian National University Printing Press.

Lynch, S.M., and J.S. Brown. 2001. Reconsidering mortality compression and deceleration: An alternative model of mortality rates. *Demography* 38(1):79–95.

Madigan, F.C. 1957. Are sex mortality differentials biologically caused? *Milbank Memorial Fund Quarterly* 35(2):202–223.

Manton, K.G., and H.D. Tolley. 1991. Rectangularization of the survival curve. *Journal of Aging and Health* 3(2):172–193.

Manuel, D.G., and J. Hockin. 2000. Recent trends in provincial life expectancy. *Canadian Journal of Public Health* 91(2):118–119.

Martel, S. 2002. *Évolution de la dispersion des durées de vie au Québec au cours du XX*[e] *siècle*. Master's thesis. Université de Montréal, Département de démographie.

Martel, S., and R. Bourbeau. 2003. Compression de la mortalité et rectangularisation de la courbe de survie au Québec au cours du XX[e] siècle. *Cahiers québécois de démographie* 32(1):43–75.

Matthews, R., and J.C. Davis. 1986. The comparative influence of regions, status, class, and ethnicity on Canadian attitudes and values, in *Regionalism in Canada*, edited by R.J. Brym. Toronto: Irwin, pp. 89–122.

McCullagh, P., and J.A. Nelder. 1989. *Generalized Linear Models.* 2nd edn. Monographs on Statistics and Applied Probability. London: Chapman & Hall.

Myers, G.C., and K.G. Manton. 1984a. Compression of mortality: Myth or reality? *The Gerontologist* 24:346–353.

————. 1984b. Recent changes in the U.S. age at death distribution: Further observations. *The Gerontologist* 24:572–575.

Nagnur, D. 1986. Rectangularization of the survival curve and entropy: The Canadian experience, 1921–1981. *Canadian Studies in Population* 13:83–102.

Nault, F. 1997. Narrowing mortality gaps, 1978 to 1995. *Health Report*s 9(1):35–41.

Nusselder, W.J., and J.P. Mackenbach. 1996. Rectangularization of the survival curve in the Netherlands, 1950–1992. *The Gerontologist* 36:773–782.

————. 1997. Rectangularization of the survival curve in the Netherlands: An analysis of underlying causes of death. *Journal of Gerontology: Social Sciences* 52B:S145–S154.

Ouellette, N. 2011. *Changements dans la répartition de décès selon l'âge : une approche non paramétrique pour l'étude de la mortalité adulte.* PhD diss. Université de Montréal, Département de démographie.

Ouellette, N., and R. Bourbeau. 2011. Changes in the age at death distribution in four low mortality countries: A nonparametric approach. *Demographic Research* 25:595–628.

Paccaud, F., C.S. Pinto, A. Marazzi, and J. Mili. 1998. Age at death and rectangularization of the survival curve: Trends in Switzerland 1969–1994. *Journal of Epidemiology and Community Health* 52(7):412–415.

Prud'homme, P. 2007. Évolution de la mortalité différentielle selon la province au Canada. Master's thesis. Université de Montréal, Département de démographie.

PHAC (Public Health Agency of Canada). 2009. Tracking Heart Disease and Stroke in Canada (Cat. No. HP32-3/2009E). Accessed via http://www.phac-aspc.gc.ca/publicat/2009/cvd-avc/pdf/cvd-avs-2009-eng.pdf (retrieved July 8, 2011).

R Development Core Team. 2010. R: A Language and Environment for Statistical Computing. Vienna, Austria: R Foundation for Statistical Computing. ISBN 3-900051-07-0.

Robine, J-M. 2001. Redéfinir les phases de la transition épidémiologique à travers l'étude de la dispersion des durées de vie : le cas de la France. *Population* 56(1-2):199–222.

Sharma, R.D., M. Michalowski, and R.B.P. Verma. 1990. Mortality differentials among immigrant populations in Canada. *International Migration* 28(4):443–450.

Statistics Canada. 2010. Deaths, 2007 (Cat. No. 84F0211X). Ottawa: Health Statistics Division.

Thatcher, A.R., S.L.K. Cheung, S. Horiuchi, and J-M. Robine. 2010. The compression of deaths above the mode. *Demographic Research* 22:505–538.

Trovato, F. 1993. Mortality differences by nativity during 1985–87. *Canadian Studies in Population* 20(2):207–223.

Trovato, F., and N.M. Lalu. 1995. The narrowing sex differential in mortality in Canada since 1971. *Canadian Studies in Population* 22(2):145–167.

————. 2001. Narrowing sex differences in life expectancy: Regional variations, 1971–1991. *Canadian Studies in Population* 28(1):89–110.

Vallin, J. 1993. Social change and mortality decline: Women's advantage achieved or regained? in *Women's Position and Demographic Change*, edited by N. Federici, K.O. Mason, and S. Sogner. Oxford: Clarendon Press, pp. 190–212.

Vaupel, J.W., W. Zhenglian, K.F. Andreev, and A.I. Yashin. 1997. *Population Data at a Glance: Shaded Contour Maps of Demographic Surfaces over Age and Time*. Monographs on Population Aging, 4. Odense: Odense University Press.

Veugelers, P.J., A.M. Yip, and D. Mo. 2001. The north-south gradient in health: Analytic applications for public health. *Canadian Journal of Public Health* 92(2):95–98.

Waldron, I. 1976. Why do women live longer than men? *Social Science and Medicine* 10:349–362.

Wilkins, R. 1980. *L'état de santé au Canada, 1926–1976*. Montréal: Institut de recherches politiques.

Wilmoth, J.R., and S. Horiuchi 1999. Rectangularization revisited: Variability of age at death within human populations. *Demography* 36(4):475–495.

Zanfongnon, R. 2008. É*volution de la mortalité différentielle selon le sexe au Canada: variations régionales, 1921–2004*. Master's thesis. Université de Montréal, Département de démographie.

Appendix A: The P-spline smoothing method

In this appendix, we first provide a brief overview of the P-spline smoothing method and then describe its use in the specific context of mortality data. For the latter task, we start by describing how the procedure works when the aim is to perform one-dimensional smoothing of mortality data over ages, as discussed by Ouellette and Bourbeau (2011). Then, we present the P-spline method in two dimensions which is used in the present paper to smooth mortality data over ages and years simultaneously.

P-splines in a nutshell

The P-spline method is a nonparametric approach that combines the concepts of B-spline and penalized likelihood. The idea behind the method is that B-splines provide flexibility, which leads to an accurate fit of the data, while the penalty, which acts on the coefficients of adjacent B-splines, ensures that the resulting fit behaves smoothly. The term *B-spline* is short for basis spline and as splines in general, B-splines are made out of polynomial pieces that are joined together at points called *knots*. The degree of the B-splines (e.g., 1: linear, 2: quadratic, 3: cubic, etc.) is given by the degree of the polynomial pieces used to build them. In this paper, we used cubic B-splines. What makes B-splines attractive is that each B-spline is nonzero on a limited range of the interval over which the smoothing procedure is taking place. This also means that in any given point of the interval, only a few B-splines are nonzero, thus offering great *local* control in the resulting fit. Increasing the number of knots in the interval will increase the amount of B-splines and enhance the ability to capture variation in the data. With the P-spline method, knots are equally spaced over the entire interval and we use a relatively large number of them, knowing that the penalty will prevent overfitting of the data by ensuring a smooth fit.

P-spline smoothing in the context of mortality data

One-dimensional smoothing

Let d_i and e_i denote respectively the observed death counts and exposure data by age i for a given year in the population under study. Also, let μ_i be the force of mortality at age i. Assuming that the force of mortality is a piecewise constant function, meaning that it is constant within each single age interval such as $\mu(x)=\mu_i$ for all $x \in [i,\ i+1)$, then $d_i \sim \text{Poisson}(e_i \cdot \mu_i)$. Thus, in order to estimate μ_i, we use a Poisson regression model such that

$$\begin{aligned} \ln(\mathrm{E}[\boldsymbol{d}]) &= \ln(\boldsymbol{e} \cdot \boldsymbol{\mu}) \\ &= \ln(\boldsymbol{e}) + \ln(\boldsymbol{\mu}) , \end{aligned}$$

where $\boldsymbol{d}$, $\boldsymbol{e}$, and $\boldsymbol{\mu}$, respectively, correspond to observed deaths, exposure, and force of mortality vectors (each one including all the age-specific information). A smooth estimate of $\boldsymbol{\mu}$ is obtained by the P-spline method (Eilers and Marx 1996):

$$\ln(\hat{\boldsymbol{\mu}}) = \boldsymbol{B}\hat{\boldsymbol{\alpha}},$$

where $\boldsymbol{B}$ is the B-spline basis regression matrix and $\hat{\boldsymbol{\alpha}}$ is the vector of estimated coefficients associated to each B-spline included in the basis. The vector of coefficients $\boldsymbol{\alpha}$ is estimated according to a maximum likelihood procedure where the penalized log-likelihood function to maximize corresponds to

$$l^* = l(\boldsymbol{\alpha},\boldsymbol{B},\boldsymbol{d}) - \tfrac{1}{2}\,\boldsymbol{\alpha}'\boldsymbol{P}\boldsymbol{\alpha}. \tag{A.1}$$

The first term on the right-hand side of this equation corresponds to the usual log-likelihood function for a generalized linear model. The second term is a penalty term ($\boldsymbol{P}$ is a penalty matrix), which forces the estimated coefficients of adjacent B-splines to vary smoothly. The trade-off between smoothness and fidelity of the model to the observed data is tuned by a smoothing parameter, which is selected according to the Bayesian Information Criterion (see Currie et al. (2004) and Camarda (2008) for further details on the penalty term in equation (A.1) and on the smoothing parameter in the context of mortality data).

Two-dimensional smoothing

To move from one- to two-dimensional Poisson P-spline smoothing, a new B-spline basis adapted for two-dimensional regression is required. Let $\boldsymbol{B}_a$ denote the B-spline basis regression matrix for ages. Similarly, let $\boldsymbol{B}_y$ be the B-spline basis regression matrix for calendar years. The new regression matrix $\boldsymbol{B}$, to be used for two-dimesional Poisson P-spline smoothing, is the following:

$$\boldsymbol{B} = \boldsymbol{B}_y \otimes \boldsymbol{B}_a, \qquad \text{(A.2)}$$

where the symbol ⊗ represents the Kronecker product.

As shown in Figure A-1, the Kronecker product of two B-splines (one along the year dimension and one along the age dimension) gives rise to a hill. Thus, the Kronecker product of the two B-spline basis, $\boldsymbol{B}_y$ and $\boldsymbol{B}_a$ in equation (A.2), will populate the age-year grid with several overlapping and regularly spaced hills such as the one of Figure A-1. Indeed, the complete illustration of $\boldsymbol{B}$ (not shown here) includes about 300 overlapping hills, which provide great flexibility in the modeling process.

Let matrices $\boldsymbol{D}$ and $\boldsymbol{E}$ denote respectively deaths and exposure data for the population under study, where rows refer to ages and columns to calendar years. In other words, matrices $\boldsymbol{D}$ and $\boldsymbol{E}$ both include as many rows as there are ages considered and as many columns as there are calendar years considered. For the purpose of regression, these data are arranged into column vectors $\boldsymbol{d} = \text{vec}(\boldsymbol{D})$ and $\boldsymbol{e} = \text{vec}(\boldsymbol{E})$; the *vec* operator *vectorizes* a given matrix by stacking its columns. Also, let $\boldsymbol{\mu}$ be the force of mortality, similarly arranged into a column vector.

Following the same idea as in the one-dimensional case described earlier, we assume that the force of mortality is constant within each single age-year interval. Thus, the Poisson regression setting, together with P-splines yield

$$\begin{aligned} \ln(\hat{\text{E}}[\boldsymbol{d}]) &= \ln(\boldsymbol{e}) + \ln(\hat{\boldsymbol{\mu}}) \\ &= \ln(\boldsymbol{e}) + \boldsymbol{B}\,\hat{\boldsymbol{\alpha}} \end{aligned} \qquad \text{(A.3)}$$

The vector of coefficients $\boldsymbol{\alpha}$ is estimated by maximizing the penalized log-likelihood function given by equation (A.1), where $\boldsymbol{B}$ is defined as in equation (A.2) and the penalty matrix $\boldsymbol{P}$ ensures that neighboring estimated coefficients vary smoothly over the age and year dimensions (see Currie et al. (2004) and Camarda (2008) for further details on the penalty matrix).

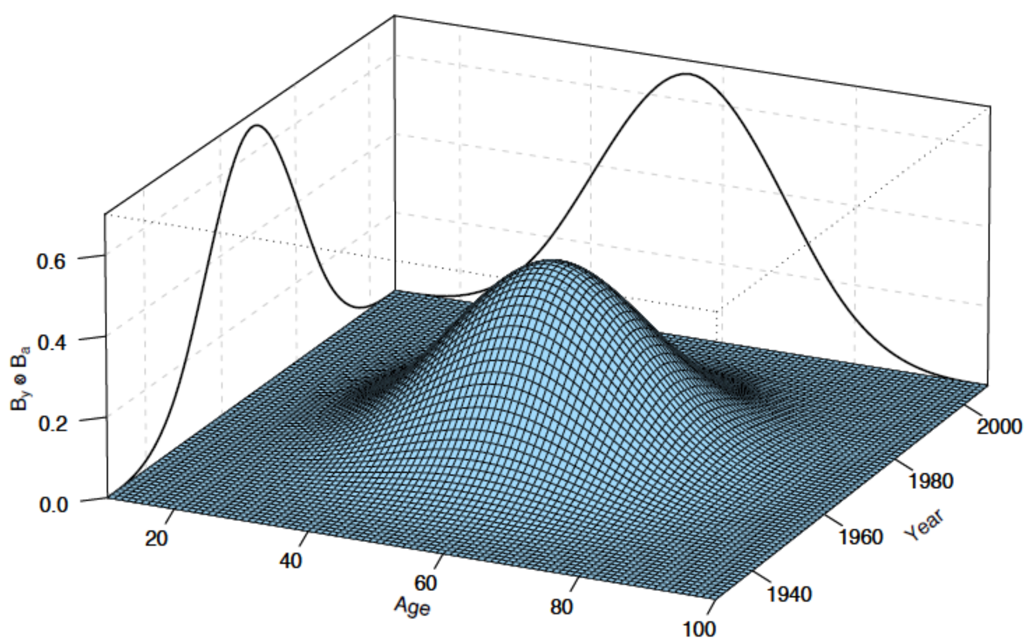

Figure A-1. Two-dimensional Kronecker product of two cubic B-splines.

The trade-off between model fidelity (or accuracy) to the actual data and model parsimony (or smoothness) is controlled by two smoothing parameters, one in each dimension. These smoothing parameters are selected independently according to the Bayesian Information Criterion (BIC), thus allowing different amount of smoothing in each dimension. Note that in order to select optimal values for the two smoothing parameters, the Akaike's Information Criterion (AIC) could also be used. However, the BIC is preferable when smoothing considerable volume of data such as mortality data with P-splines because it penalizes model complexity more heavily than AIC. Random fluctuations in the final fitted mortality surface are thus less likely to be found when BIC rather than AIC is used to select the smoothing parameters (see Currie et al. (2004) and Camarda (2008) for further details on the selection of the smoothing parameters in the context of mortality data).

More explicitly, from equations (A.2) and (A.3), the smoothed force of mortality corresponds to:

$$\begin{aligned}\hat{\boldsymbol{\mu}} &= \exp(\boldsymbol{B}\hat{\boldsymbol{\alpha}}) \\ &= \exp\left((\boldsymbol{B}_y \otimes \boldsymbol{B}_a)\hat{\boldsymbol{\alpha}}\right). \end{aligned} \tag{A.4}$$

Appendix B: 99% bootstrap confidence intervals

In this appendix, we present a brief outline of the bootstrap approach used in this paper. We also provide an excerpt of the bootstrap confidence intervals results for the modal age at death estimates (see Table B-1) and standard deviation above the mode estimates (see Table B-2). Further details on bootstrap methods in general can be found in Efron and Tibshirani (1993). Moreover, Koissi et al. (2006) recently developed a residual bootstrap approach to build confidence intervals for forecasted life expectancy as an extension to the original Lee-Carter method. We adapted their approach to the P-splines mortality modeling.

In the present study, the residual bootstrap method allows us to estimate the variability of the modal age at death estimates and the standard deviation above the mode estimates for which theoretical distributions are unknown. From the fitted force of mortality $\widehat{\boldsymbol{\mu}}$ given in equation (A.4), one can obtain the matrix of fitted death counts $\widehat{\boldsymbol{D}}$. Indeed, from equations (A.3) and (A.4), we have $\widehat{\boldsymbol{d}} = \boldsymbol{e} \cdot \widehat{\boldsymbol{\mu}}$, where $\widehat{\boldsymbol{d}} = \text{vec}(\widehat{\boldsymbol{D}})$. Deviance residuals (McCullagh and Nelder 1989: 39-40) can thus be computed since

$$\boldsymbol{r} = \text{sign}\big(\boldsymbol{D} - \widehat{\boldsymbol{D}}\big) \cdot \sqrt{2\left[\boldsymbol{D}\ln\left(\frac{\boldsymbol{D}}{\widehat{\boldsymbol{D}}}\right) - \boldsymbol{D} + \widehat{\boldsymbol{D}}\right]}\,. \tag{B.1}$$

From these deviance residuals, we sample with replacement an entire new set of residuals $\boldsymbol{r}_b^*$ called the bootstrapped residuals. Replacing deviance residuals $\boldsymbol{r}$ by bootstrapped residuals $\boldsymbol{r}_b^*$ in equation (B.1) and rearranging this equation leads to

$$\widehat{\boldsymbol{D}} - \boldsymbol{D}\ln\big(\widehat{\boldsymbol{D}}\big) + \boldsymbol{r}_b^{*2} + \boldsymbol{D} - \boldsymbol{D}\ln(\boldsymbol{D}) = 0. \tag{B.2}$$

Equation (B.2) can be solved numerically in order to obtain a new matrix of bootstrapped deaths $\widehat{\boldsymbol{D}}_b^*$. Together with the matrix of exposure data $\boldsymbol{E}$, these bootstrapped deaths are then used in a two-dimensional P-spline framework (see Appendix A) and lead to new bootstrapped coefficients $\widehat{\boldsymbol{\alpha}}_b^*$. Corresponding smoothed bootstrapped density function describing the two-dimensional age-at-death distributions and bootstrapped modal age at death estimates are then computed.

The procedure described above, starting with the residual sampling with replacement step, was repeated 5,000 times. This led to a bootstrapped distribution of $\widehat{M}(y)$ and $\widehat{SD(M+)}(y)$, from which 0.995th and 0.005th empirical percentiles were extracted and used as lower and upper confidence bounds for the 99% bootstrap confidence intervals. As an example, Figure B-1 shows the female modal age at death estimates along with their 99% confidence intervals for British Columbia and Quebec.

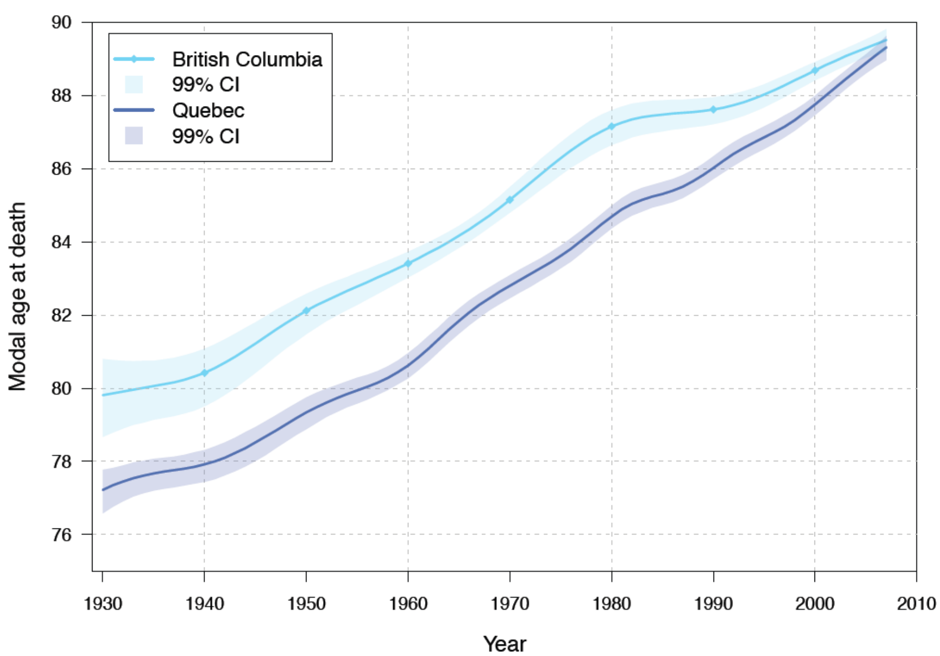

Figure B-1. Modal age at death estimates and corresponding 99% bootstrap confidence intervals, British Columbia and Quebec, 1930–2007.

Table B-1. Modal age at death estimates and corresponding 99% bootstrap confidence intervals by province and sex, Canada, 1930, 1950, 1970, 1990, and 2007 (Females); 1970, 1980, 1990, 2000, and 2007 (Males).

Sex	Year	BC	AB	SK	MB	ON	QC	NB	NS	PE	NL
Females	1930	79.8	78.7	79.5	79.2	78.5	77.2	78.6	79.3	82.1	-
		(78.7,80.8)	(76.9,80.3)	(78.9,80.1)	(78.5,79.9)	(78.0,79.0)	(76.6,77.8)	(77.9,79.1)	(78.3,80.1)	(80.8,82.5)	-
	1950	82.1	81.4	81.0	81.2	81.2	79.3	80.2	81.4	82.6	80.9
		(81.5,82.6)	(80.5,82.1)	(80.6,81.3)	(80.7,81.5)	(80.4,81.3)	(78.9,79.8)	(79.7,80.5)	(80.9,81.8)	(81.7,82.8)	(80.2,81.6)
	1970	85.2	84.8	85.1	84.4	84.0	82.8	83.7	83.9	85.1	83.4
		(84.8,85.5)	(84.4,85.2)	(84.8,85.4)	(84.0,84.8)	(83.8,84.2)	(82.5,83.1)	(83.3,84.0)	(83.5,84.2)	(84.3,85.3)	(82.9,83.8)
	1990	87.6	88.1	88.1	87.2	86.9	86.0	86.8	86.5	87.2	85.1
		(87.2,88.0)	(87.7,88.5)	(87.8,88.4)	(86.8,87.6)	(86.6,87.1)	(85.7,86.3)	(86.4,87.1)	(86.0,86.9)	(86.5,87.4)	(84.7,85.6)
	2007	89.5	89.7	89.7	89.5	89.4	89.3	88.5	88.3	88.5	86.9
		(89.2,89.8)	(89.3,90.2)	(89.4,90.0)	(89.1,89.8)	(89.2,89.6)	(89.0,89.6)	(88.1,88.9)	(87.9,88.7)	(87.7,88.9)	(86.3,87.5)
Males	1970	78.5	78.6	80.5	78.4	76.3	77.2	78.9	77.2	79.2	79.6
		(76.7,80.0)	(77.4,79.6)	(79.6,81.5)	(77.2,79.2)	(75.7,77.1)	(76.0,78.0)	(77.4,79.7)	(75.9,78.4)	(77.7,79.6)	(77.9,80.6)
	1980	80.7	79.8	81.3	79.4	78.3	77.7	79.1	77.4	79.5	79.0
		(80.0,81.5)	(78.8,80.8)	(80.3,82.1)	(78.4,80.3)	(77.7,78.8)	(76.7,78.3)	(78.3,79.7)	(76.3,78.3)	(78.8,80.4)	(77.7,79.9)
	1990	82.2	82.1	82.1	81.4	80.9	79.9	80.0	79.4	80.5	79.6
		(81.6,82.8)	(81.2,82.8)	(81.0,82.8)	(80.5,82.2)	(80.5,81.3)	(79.3,80.6)	(79.4,80.6)	(78.5,80.2)	(79.3,80.9)	(78.8,80.3)
	2000	84.6	84.1	83.4	83.1	83.5	82.3	82.2	82.0	82.1	81.0
		(84.2,85.0)	(83.6,84.6)	(82.4,84.1)	(82.3,83.6)	(83.2,83.8)	(81.7,82.8)	(81.1,83.1)	(81.1,82.8)	(81.4,82.8)	(80.2,81.7)
	2007	86.9	85.9	85.5	84.3	85.7	85.7	84.6	83.9	83.3	82.3
		(86.1,87.6)	(85.0,86.9)	(84.5,86.4)	(83.4,85.0)	(85.3,86.3)	(84.8,86.6)	(83.5,85.6)	(82.6,84.9)	(82.1,83.9)	(81.2,83.1)

Note: 99% bootstrap confidence intervals results are based on 5,000 bootstrap samples.

Table B-2. Standard deviation above the modal age at death estimates and corresponding 99% bootstrap confidence intervals by province and sex, Canada, 1930, 1950, 1970, 1990, and 2007 (Females); 1970, 1980, 1990, 2000, and 2007 (Males).

Sex	Year	BC	AB	SK	MB	ON	QC	NB	NS	PE	NL
Females	1930	8.44	8.04	7.99	8.30	7.58	8.07	7.81	8.26	7.10	-
		(7.49,8.83)	(6.98,8.75)	(7.35,8.16)	(7.60,8.53)	(7.23,7.77)	(7.69,8.39)	(7.38,8.06)	(7.77,8.68)	(6.56,7.46)	-
	1950	7.31	7.46	7.82	7.49	7.23	7.72	7.87	7.68	7.21	7.59
		(6.94,7.50)	(6.96,7.73)	(7.46,7.92)	(7.13,7.63)	(7.14,7.56)	(7.47,7.92)	(7.55,8.02)	(7.37,7.86)	(6.83,7.43)	(7.04,7.91)
	1970	7.14	7.36	7.23	7.29	7.31	7.09	7.45	7.49	7.32	7.24
		(6.86,7.25)	(7.02,7.46)	(6.92,7.32)	(6.97,7.44)	(7.20,7.42)	(6.90,7.24)	(7.14,7.58)	(7.23,7.63)	(6.92,7.54)	(6.86,7.40)
	1990	6.96	6.72	6.70	7.19	7.07	7.44	7.30	7.57	7.18	7.57
		(6.74,7.12)	(6.45,6.88)	(6.45,6.81)	(6.91,7.37)	(6.92,7.16)	(7.23,7.57)	(7.00,7.43)	(7.28,7.74)	(6.82,7.42)	(7.20,7.76)
	2007	6.51	6.41	6.08	6.35	6.42	6.71	6.86	6.91	6.55	6.90
		(6.26,6.70)	(6.11,6.63)	(5.80,6.25)	(6.05,6.56)	(6.25,6.51)	(6.47,6.92)	(6.52,7.09)	(6.58,7.15)	(6.11,6.93)	(6.48,7.23)
Males	1970	8.78	8.81	7.97	8.65	9.15	8.44	8.32	8.99	8.55	7.98
		(8.02,9.59)	(8.29,9.36)	(7.47,8.35)	(8.20,9.16)	(8.79,9.46)	(8.01,8.99)	(7.83,8.95)	(8.38,9.56)	(8.07,8.94)	(7.37,8.62)
	1980	8.33	8.58	8.01	8.54	8.84	8.90	8.51	9.10	8.63	8.38
		(7.92,8.60)	(8.05,9.02)	(7.55,8.43)	(8.05,8.99)	(8.57,9.07)	(8.54,9.35)	(8.14,8.82)	(8.60,9.57)	(8.15,8.98)	(7.83,8.88)
	1990	7.65	7.67	7.87	7.90	7.93	8.14	8.32	8.40	8.36	8.13
		(7.35,7.94)	(7.30,8.05)	(7.45,8.29)	(7.48,8.25)	(7.70,8.08)	(7.81,8.40)	(7.91,8.52)	(7.95,8.80)	(7.91,8.68)	(7.68,8.42)
	2000	6.94	6.99	7.42	7.53	7.13	7.45	7.62	7.67	7.81	7.52
		(6.70,7.13)	(6.73,7.20)	(7.03,7.89)	(7.20,7.82)	(7.02,7.29)	(7.21,7.71)	(7.12,8.09)	(7.21,8.06)	(7.33,8.18)	(7.09,7.82)
	2007	6.56	6.56	6.76	7.36	6.90	6.82	6.93	7.27	7.40	6.97
		(6.16,6.94)	(6.03,7.00)	(6.21,7.19)	(6.90,7.74)	(6.55,7.07)	(6.34,7.22)	(6.32,7.44)	(6.70,7.87)	(6.85,7.86)	(6.46,7.43)

Note: 99% bootstrap confidence intervals results are based on 5,000 bootstrap samples.