Found in translation? A cohort perspective on tempo-adjusted life expectancy *

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Summary. What does tempo-adjusted period life expectancy measure? Taking a cohort perspective, I show that under conditions of constant linear mortality shifts the tempo-adjusted period indicator translates exactly to the cohort born $e_0^*(t)$ years earlier. I discuss the implications of cohort translation for the interpretation and application of tempo-adjusted period life expectancy.

1 Introduction

Life expectancy at birth is at root a cohort concept. It tells us how long, on average, the members of a cohort survive. Actual life expectancy can only be known fully for cohorts born long ago. To summarize recent mortality conditions and period-to-period variation, the hypothetical concept of period life expectancy is conventionally used. But even period life expectancy refers conceptually to a cohort – the hypothetical one that lives according to the rates observed in a single period.

When mortality conditions are improving, period life expectancy is *less* than that of the cohort born in the period. This is because the hypothetical cohort following the period life table is deprived of future mortality improvement.

I recite this basic property of period life expectancy because the "tempo adjusted" method of measuring period life expectancy – as developed by Bongaarts & Feeney (2002) – arrives at exactly the opposite conclusion. According to Bongaarts and Feeney, period life expectancy *overstates* longevity when mortality conditions are improving. They conclude: "Our main finding is that the conventional calculation of period life expectancy at birth gives a misleading indication of how long we live. We are not living as long as we thought we were." (p. 25).

To be fair, Bongaarts and Feeney, except at a few points, are not talking about cohorts. Instead, they intend e_0^* as a period measure that tries to improve upon period life expectancy. What such an improved period indicator actually measures is the subject of much debate as is clear from many of the chapters in this volume. The approach taken here is to recast tempo-adjustment in cohort terms. Doing this enables us to resolve the counter-intuitive direction of tempo-adjustment by showing which cohort B&F are referring to when they say "we."

The approach is similar to that of Goldstein & Wachter (2004), which showed – using a different model of temporal mortality change – the correspondence between period life expectancy e_0 and the life expectancy of particular cohort. Here, I look at which cohort has the life expectancy equal to current *tempo-adjusted* life expectancy e_0^*. I find that under linearly shifting mortality, defined below, tempo-adjusted life expectancy for year t translates to the cohort *dying* in year t: this is the cohort *born* e_0^* years earlier.[2]

An additional assumption is needed for this simple cohort translation of e_0^* to hold exactly. B&F's tempo-adjustment assumes that deaths are postponed uniformly across all ages, with the size of the shift possibly varying from year to year. To this I add the assumption that the size of the shift is constant from year to year, a pattern I call "linear shifts." As will be seen, the linear shift pattern is consistent with quite recent mortality trends above age 30 in low mortality populations. The linear shift assumption is, however, not a general feature of human populations. Prior to World War II, change was distinctly non-linear in many countries. It remains to be seen whether the recent linear shift pattern will continue.

Under the linear shift model, the current tempo-adjusted period life expectancy has the same value as the life expectancy of a past cohort. This correspondance with cohorts from the past explains why Bongaarts and Feeney's measure is less, not more, than current period life expectancy.

Furthermore, under linear shifts, it is possible to obtain directly the life expectancy of the cohort born in every period, including the current one, a quantity that is arguably of more interest than e_0^*.

Neither B&F's tempo-adjustment nor the discussion presented here applies to life expectancy at birth. Instead, both ignore all mortality before about age 30. For notational simplicity, the current discussion follows B&F, using the shorthand of e_0, e_0^*, and e_0^c to refer to the period, tempo-adjusted period, and cohort life expectancies at birth, assuming no mortality below age 30. In traditional demographic notation, these quantities would be written $e_{30} + 30$, $e_{30}^* + 30$, and $e_{30}^c + 30$.[3]

[2] This result was suggested in simulation by Bongaarts (2004), who also found that it held approximately in modern real-world populations.

[3] Although we use B&F's shorthand here, it is worth keeping in mind that although mortality below age 30 is low in modern industrialized populations, $e_{30} + 30$ does not equal e_0. In the 2002 Swedish female period life table, ignoring under-30

2 Proof of exact cohort translation

Let $l^c(a,t)$ be the surviving proportion of a cohort born at time $t-a$ and aged a at time t. For all $a \leq 0$, define $l^c(a,t) = 1$ for all t. This formulation amounts to the same thing as B&F's requirement of no mortality below age 30.

A proportionally shifting surface $l^c(a,t)$ consistent with B&F's proportionality assumption is obtained by shifting the baseline $l^c(a,0)$ up or down the age axis by an amount $F(t)$ such that

$$l^c(a,t) = l^c(a - F(t), 0), \tag{1}$$

again letting $l^c(a,t) = 1$ for $a - F(t) \leq 0$. The fact that $F(t)$ is not a function of age is the B&F's proportionality assumption. The additional assumption of linearity over time in the shifts can be introduced by letting $F(t) = rt$.

The cohort born at time τ has life expectancy

$$e_0^c(\tau) = \int_0^\infty l^c(a, \tau + a)\, da.$$

Following Bongaarts & Feeney (in this volume p. 11), the adjusted period life expectancy $e_0^*(t)$ is equal to

$$CAL(t) = \int_0^\infty l^c(a,t)\, da.$$

I use the CAL notation to emphasize its correspondence with the "cross sectional average length of life" introduced by Brouard (1986) and developed by Guillot (2003) .[4]

I want to show that

$$e_0^c(\tau) = CAL(\tau + e_0^c(\tau)). \tag{2}$$

Showing this demonstrates that the approximation given by Bongaarts (2004),

$$e_0^c(t - e_0^*(t)) \approx e_0^*(t), \tag{3}$$

actually holds exactly.[5]

This equality is shown as follows by expressing $e_0^c(\tau)$ and $CAL(\tau + e_0^c(\tau))$ in terms of $CAL(0) = \int_0^\infty l^c(a,0)\, da$.

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mortality increases life expectancy by 0.6 years, more than a third of the 1.6 years tempo-effect that B&F find for Sweden 1980-1995.

[4] The quantity CAL used here differs from that used by Brouard and Guillot in that it assumes no child or young adult mortality under a given age such as 30.

[5] To see the correspondance, substitute $t = \tau + e_0^c(\tau)$ and note that from (2) $e_0^c(\tau) = CAL(t) = e_0^*(t)$.

Linear proportional shifts means that the cohort born at time τ has a survival curve that resembles the initial profile $l^c(a, 0)$, except that each age "a" is shifted to age $a - r(\tau + a)$. In effect, a member of the cohort "feels younger" than they are by a factor of $r(\tau + a)$, where the $r\tau$ term accounts for the improvements up to the date at which the cohort is born, and the ra term accounts for the additional improvements obtained by the time the cohort reaches age a. Cohort period life expectancy under linear shifts can be written in terms of the baseline survival at time 0 as

$$e_0^c(\tau) = \int_0^\infty l^c(a - r(\tau + a), 0)\, da.$$

To evaluate, substitute $u = a(1 - r) - r\tau$ and $da = du/(1 - r)$. This gives

$$e_0^c(\tau) = \frac{1}{1 - r} \int_{0 - r\tau}^{\infty} l^c(u, 0)\, du. \tag{4}$$

Recalling that for $u \leq 0$, $l^c(u) = 1$, the integral evaluates to

$$e_0^c(\tau) = \frac{CAL(0) + r\tau}{1 - r}. \tag{5}$$

We can evaluate $CAL(\tau + e_0^c(\tau))$ in a similar manner. The linearly shifting age distribution means that $CAL(t)$ is simply growing linearly with time. For any t,

$$CAL(t) = \int_0^\infty l^c(a, t)\, da = \int_0^\infty l^c(a - rt, 0)\, da.$$

Substituting $u = a - rt$ and $du = da$,

$$CAL(t) = \int_{0 - rt}^{\infty - rt} l^c(u)\, du = CAL(0) + rt. \tag{6}$$

To prove (2), we are interested in $t = \tau + e_0^c(\tau)$. From (6),

$$CAL(\tau + e_0^c(\tau)) = CAL(0) + r(\tau + e_0^c(\tau)).$$

Substituting from (5) for $e_0^c(\tau)$,

$$CAL(\tau + e_0^c(\tau)) = CAL(0) + r\tau + r\frac{CAL(0) + r\tau}{1 - r},$$

which simplifies to

$$CAL(\tau + e_0^c(\tau)) = \frac{CAL(0) + r\tau}{1 - r}. \tag{7}$$

The right-hand side of this last expression is identical to the right-hand side of equation (5) for $e_0^c(\tau)$, which is what we wanted to show to prove (2).

Note that this change of variable approach is perfectly general for any survival curve $l^c(a, 0)$ and any $r \neq 1$. It does not require Gompertzian survival or any other particular form of the hazards.

3 Discussion

We have shown that tempo-adjusted period life expectancy $e_0^*(t)$ under linear shifts is equal to the life expectancy of the cohort dying in that year t.

The equality of tempo-adjusted life expectancy with lagged cohort life expectancy provides us with an alternative way to think about e_0^*. Whereas B&F use e_0^* as a counterfactual estimate of period mortality corrected for tempo distortion, we have shown here that in the context of steadily shifting survival curves e_0^* is also a measure of cohort life expectancy.

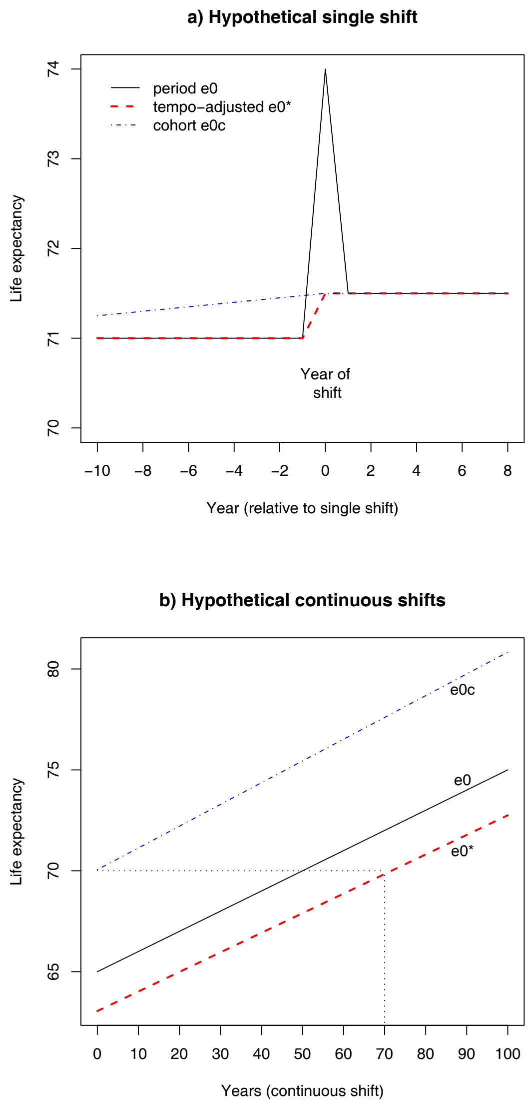
Both interpretations are interesting and potentially useful. I would argue that the B&F interpretation is most valuable in conditions of sudden mortality change, whereas the cohort interpretation is more valuable in conditions of steady mortality change.

Below I lay out two extreme scenarios that help us to understand the difference. Bongaarts and Feeney introduced the first in their 2002 paper; the second is explored by their paper in this volume as well as the chapter by Rodriguez (in this volume).

3.1 A single magic pill

The story of the "life extension" pill discussed by Bongaarts and Feeney (in this volume p. 11) illustrates the potential advantages of tempo-adjustment in the case of a sudden shift in survival. (See Figure 1a.) On January 1, everyone in a previously stationary population takes a pill postponing their previously programmed date of death by 3 months. Everyone born afterwards also takes the pill. The effect of such a pill in the year it is taken is to reduce the number of deaths by one-fourth, since no one will die in the first 3 months of the year. In the year the pill is introduced, death rates also fall by about one-fourth, raising life expectancy dramatically, not by the three months indicated by the pill but by several years because of the enormous drop in death rates.

In the case of a single such pill, period life expectancy spikes in the year the pill is taken and then falls thereafter to a constant value equal to the pre-pill life expectancy plus the extension granted by the pill. This makes the measurement taken in the year the pill appeared suspect, a candidate for tempo-adjustment. As panel (a) shows, e_0 shows a spike in the year the pill is introduced, but e_0^* shows no spike, instead attributing the appropriate 3 month increase in life expectancy.



Source: Panel (a) from B&F (2003) Figure 5, with illustrative $e_0^c(t)$ added. Panel (b) calculated for e_0 growing linearly from 65 to 75 at a rate of 0.1 years per year, with $e_0^*(t) = e_0(t)[1 - r * H]$ and $e_0^c(t) = e_0^*/(1 - r)$, where $r = 0.1$ and $H = .3$.

Fig. 1. Time paths of period and cohort life expectancy and of tempo-adjusted period life expectancy in (a) single shift scenario and (b) linear shift scenario.

The figure also shows cohort life expectancy for those born in each year. From a cohort view, adjusted-life expectancy performs well in the year that the pill is taken. In that year, unadjusted e_0 overestimates the life expectancy of the cohort being born, but the adjusted period measure e_0^* accurately predicts cohort life expectancy. In the years following the pill introduction, both e_0^* and e_0 are equal to e_0^c. In the years before the pill is taken, however, neither period life expectancy nor adjusted period life expectancy matches cohort life expectancy because neither period measure can foresee the subsequent sudden increase in longevity.

The lesson to be drawn from this scenario is that under a sudden mortality shock, akin to the one-time pill[6], e_0^* provides a better indication of the implications of the shock for cohort mortality than does e_0.

3.2 A series of magic pills

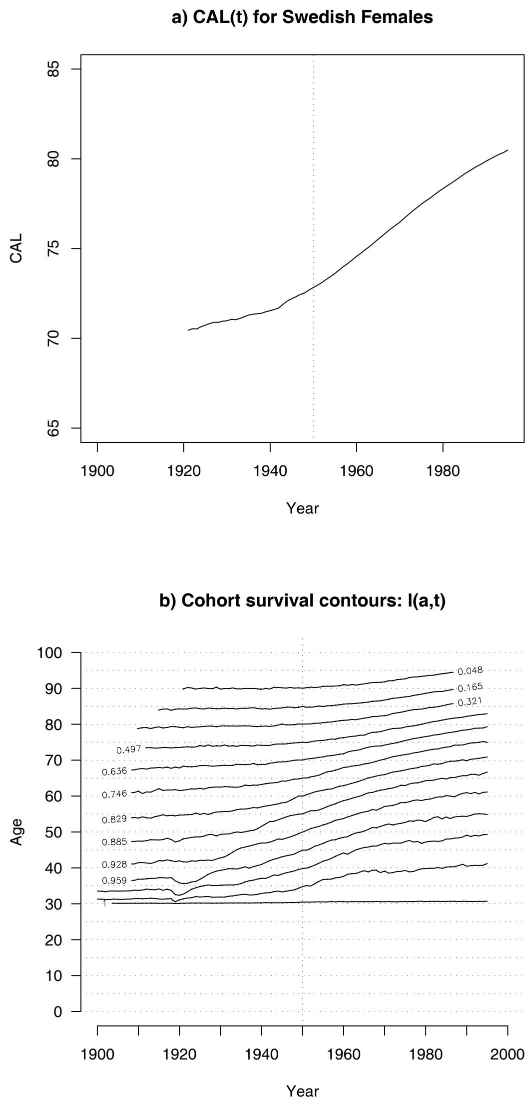
Now let us consider the case where such pills are given year after year, continually re-extending life by some constant amount each year. This scenario is the one investigated mathematically above and is illustrated in Figure 1b. In this case, we still have $e(t)$ larger than $e_0^*(t)$, but rather than this difference occurring in a single year as in the single-pill example, it persists over time. Now, the equality of $e_0^*(t)$ is not with the cohort born in year t, as in the single year example, but rather with the cohort born $e_0^*(t)$ years earlier that is *dying* in year t. This result from the formal analysis is illustrated by the dotted lines in the figure showing that $e_0^c(0) = 70 = e_0^*(70)$.

Under linear shifts that result from a series of pills, it would clearly be wrong to interpret tempo-adjusted life expectancy as an estimate of the cohort born in year t. The adjustment moves period life expectancy farther from, not closer to, that of the cohort.

3.3 Which scenario is more realistic?

We can now ask which of these two cases bears more resemblance to observed patterns of mortality change. Here there is no debate. Bongaarts and Feeney (2002 and in this volume p. 11) answer this question quite clearly in their empirical analysis of 1980-1995 France, Sweden and the United States (See for example figure 6 of B&F in this volume p. 11). In every case, the improvement of mortality by all measures has been steady. There is no historical example that bears any resemblance to the one-time pill example. Tuljapurkar, Li & Boe (2000) show, using methods different from the shift model, that since World War II, steady mortality decline is the rule throughout the industrialized world.

[6] See Le Bras (in this volume), who argues that even those one-time shocks that are observed do not occur in a manner that delay or advance deaths uniformly by age.



Note: Both panels assumes no mortality under age 30. $CAL(t) = \int_{30}^{\infty} l^c(a,t)\, da + 30$. Since $r = \frac{d}{dt} CAL(t)$, constant shifts imply linear $CAL(t)$. Source: Calculated from $m(x,t)$ data from www.mortality.org.

Fig. 2. Observed time paths of mortality change for Swedish females. Panel (a): Mortality change over all ages as measured by $\mathbf{CAL(t)}$ Panel (b): Contour plot of cohort survival $\mathbf{l^c(a,t)}$ using isoclines intersecting $\mathbf{l^c(30)}, \mathbf{l^c(35)}, \ldots, \mathbf{l^c(95)}$ in 1950.

It is useful to look at a longer course of time. Panel (a) shows CAL – what Bongaarts and Feeney call e_0^* – for Sweden females from 1920 to 1995. Bongaarts & Feeney (in this volume p. 11) figure 6 shows the last 15 years of this series. Linearity in CAL implies linear shifts. We see that the near linearity they find for 1980-1995 is a continuation of the post-World War II pattern. Before this, however, the pace of improvement was considerably slower. There is no evidence from looking at CAL of sporadic large mortality shifts of the kind in the single-pill scenario. Rather, the last half-century has been consistent with the linear shift scenario.

We can see in detail at how close both the proportionality and linear shift assumptions hold by looking at the full $l^c(a,t)$ surface (Panel (b)). The contour plot shows the isoclines of $l^c(x,t)$ at the levels seen in 1950 for $x = 30, 35, \ldots 95$. For example, the contour labeled "0.048" shows the age at which $l^c(90, 1950)$ is reached over the course of the century, and we can see that by 1995 this level of survival was reached at age 95 rather than 90.

Proportionality can be checked by looking at whether the slopes at different ages change simultaneously. The linearity of the shifts requires further that the contours be straight lines. The figure shows there were few shifts at all in the first two decades of the century in Sweden. Starting after World War I, and the influenza epidemic, survival to younger ages started to shift, followed by shifts in survival to older ages after World War II. Since about 1950, the contours are nearly linear and nearly parallel, particularly above age 60, when most deaths are occurring. Overall, neither proportionality or linearity seems a good description for the whole century. However, the linear shift model does not seem at odds with recent decades. The only evidence of mortality change that resembles the single-pill example is perhaps the 1918 influenza epidemic, but even this does not appear across all ages.

3.4 Telling the future

If we expect linear shifts well into the future, then we can go one step further. We have seen that under linear shifts, e_0^* understates even more dramatically than period life expectancy the survival of those born in a period. However, the same derivation we used to show the cohort that has life expectancy $e_0^*(t)$ can also be used to show the life expectancy of the cohort born in year t. Replacing τ with t and substituting from (6), we find

$$e_0^c(t) = e_0^{**} = \frac{e_0^*(t)}{1-r}, \tag{8}$$

where we use e_0^{**} to denote the rescaled e_0^*. Although $e_0^*(t)$ is itself a rather out-of-date measure referring to a cohort born long before t, the simplicity of the linear shift model allows us to go from $e_0^*(t)$ to the cohort born in year t by rescaling.

These exact relationships for steady mortality change should hold approximately when there are small variations in the pace mortality improvement.

If there is no temporal autocorrelation, the variations will cancel each other out. Such random ups and downs seem to encompasses the modern experience of mortality decline in advanced industrial countries, forming the basis of the Lee-Carter stochastic forecasting method (Lee and Carter 1992). Systematic slowdowns or accelerations that last many years can make the relationship between e_0^* and e_0^c quite different from the results found here. [7]

3.5 The order of mortality measures

With an exact expression for cohort life expectancy, we can now provide a full description of the ordering of different measures of life expectancy and their cohort translations under linear shifts. Table 1 shows tempo-adjusted period life expectancy, unadjusted period life expectancy, and rescaled tempo-adjusted life expectancy for Sweden using the same data as B&F and the cohort translation of these quantities. The table reiterates the point we began with that cohort life expectancy is larger, not smaller, than period life expectancy if we are considering the cohort born in the period. It shows that e_0^* actually refers to the cohorts born around 1900-1915, not cohorts born in 1980-1995.

In this example, period life expectancy and tempo-adjusted period life expectancy are close to each other relative to cohort life expectancy. The ordering $e_0^* < e_0 << e_0^{**}$ applies quite generally in conditions of improving mortality. Letting H denote Keyfitz's measure of life table entropy, e_0^* increases the observed mortality rates by a factor of about $1 + r$, which reduces life expectancy by about a factor of $1 - Hr$ (Keyfitz 1985). Life table entropy is small, on the order of 0.2, and so if $r = 0.1$, $e_0^* \approx 0.98e_0$. To see that e_0^{**} is larger than either of these, note that dividing e_0^* by $1 - r$ gives a quantity substantially greater than e_0.[8]

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[7] A more general expression for cohort mortality can be given as follows. Let r_t be the shift in year t and R_t be the cumulative shift $\int_0^t r_t dt$. In this case, cohort life expectancy is given in terms of the baseline survival profile as

$$e_0^c(\tau) = \int_0^\infty l^c(a - R_{\tau+a}, 0)\, da.$$

Substituting $u = a - F(\tau + a)$ and $da = du/[1 - r(\tau + a)]$,

$$e_0^c(\tau) = \int_{-F_\tau}^\infty l^c(u)[1 - r_{\tau+a}]^{-1}\, du.$$

This reduces to (4) for when $r_{\tau+a}$ is a constant. When $r_{\tau+a}$ varies only slightly and in a manner that is uncorrelated with $l^c(u)$, then fluctuations should not influence $e_0^c(\tau)$ much, since shifts larger-than-average shifts will be cancelled out by smaller-than-average shifts.

[8] Formally, if entropy is sufficiently large then the inequality need not hold. But high H implies a high variance of age at death, typically in the form of high mortality among children, an age-group that is excluded from the Bongaarts-Feeney shift model.

Table 1. Ordering and cohort translation of period and tempo-adjusted period measures under linear shifts.

Period or tempo-adjusted period measure	Estimate for Sweden, 1980-95	Cohort translation	
$e^*_{30} + 30$	79.4	$e^c_0(t - e^*_0)$	$\approx e^c_0(1900 - 1915)^a$
$e_{30} + 30$	81.1	$e^c_0(t - \lambda)$	$\approx e^c_0(1905 - 1920)^b$
$e^{**}_{30} + 30 = \frac{e^*_0+30}{(1-r)}$	94.5	$e^c_0(t)$	$\approx e^c_0(1980 - 1995)^c$

[a] This cohort life table value was reached by the cohort of 1909, according to www.mortality.org.
[b] $\lambda < e^*_0$ but exact value unknown; 1905-1920 is a rough estimate.
[c] Assuming continued linear shifts.

Values for $e^*_{30} + 30$ and $e_{30} + 30$ from B&F(2004) Table 1. $e^{**}_{30} + 30$ calculated as $\frac{e^*_0}{(1-r)}$ using $r = 0.16$ as estimated by author from www.mortality.org. In "Cohort translation" column e^c_0 is used as shorthand for $e^c_{30} + 30$.

Without a crystal ball, we don't know for sure how long the cohorts born from 1980 to 1995 will live. But what we do know, assuming continued mortality decline, is that e^*_0 is clearly the worst measure, giving an even lower figure than the already too low period life expectancy. If we are going to adjust period life expectancy, we should readjust it again to produce not the cohort born long ago, but rather our best guess at the cohort born today, e^{**}.[9]

4 Conclusion

Some critics of Bongaarts and Feeney's theory of mortality tempo effects argue that its assumption of uniform postponement of death across all ages is unrealistic. Others argue that e^*_0 is not really a period measure, but rather depends on the history of the population. In this chapter, my approach has not been to try to debunk tempo-adjustment but rather to take it even further by assuming that the Bongaarts and Feeney's uniform shift repeats itself over many decades – so long that cohort mortality becomes a simple function of the baseline mortality schedule and the pace of the shift.

Under these conditions, two results were found. First, $e^*_0(t)$ translates to cohort life expectancy for those born $e^*_0(t)$ years earlier, long before the period under consideration. Second, the cohort life expectancy of those born today, or in any year t, can be found by a simple inflation of $e^*_0(t)$. Viewed this way, $e^*_0(t)$ itself is not a measure of great interest. It does not tell us what is

[9] Incidentally, this figure of 94.5 years is not out-of-line with optimistic forecasts. Oeppen & Vaupel (2002) predict that record period life expectancy will be 95 by 2040, which would apply to cohorts born about 1970 or 1980 (Goldstein and Wachter 2004).

happening in year t – this is given by the unadjusted period life table. It does not tell us the future – this is given by the life table of the new-born cohort. Rather it tells us about the cohort born in the past that is, on average, dying in year t.[10]

If mortality change were to be sudden, and to occur in such a way as to advance or to postpone deaths uniformly across all ages, tempo-adjustment could produce measures giving a valuable sense of the implications of the mortality rates seen during shocks. The difficulty, so far, is that mortality change has not occurred in this way. Recent history in the industrialized world has been has been one of steady, not sudden, mortality change. In this context, the linear shift model provides a framework for understanding what tempo-adjusted life expectancy is actually measuring and for developing even more informative indicators.

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[10] The backward-looking tendency of e_0^* is not due to the way it is measured, which after all is from current data, but rather from fact that the post-adjustment longevity estimates are equal to those of cohorts born long ago. Wachter (in this volume) shows that the differential equations that define e_0^* effectively define a moving average of recent period life expectancies.

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