Competing effects on the average age of infant death

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Abstract

In recent decades, the relationship between the average length of life for those who die in the first year of life – the lifetable quantity ${}_1a_0$ – and the level of infant mortality, on which its calculation is often based, has broken down. The very low levels of infant mortality in the developed world correspond to a range of ${}_1a_0$ quantities. We illustrate the competing effect of falling mortality and reduction in preterm births on ${}_1a_0$, through simulation and through an example with US data on two populations with very different levels of premature birth – infants born to non-Hispanic white mothers and to non-Hispanic black mothers. We demonstrate that falling mortality reduces ${}_1a_0$ while fewer premature births increase it. Future work will focus on utilizing these observations in developing an alternative formula for ${}_1a_0$, which aims to capture not only the level of infant mortality, but also the level of prematurity.

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1 Introduction

The average length of life for those who die in the first year of life, ${}_1a_0$, is an important lifetable quantity, the first building block of the calculation of person-years lived that ultimately sums to the expectation of life at birth. However, the data required to calculate ${}_1a_0$ exactly are often not readily available, and as such, producers of lifetables usually rely on empirical relationships between the overall level of infant mortality and average age at death to calculate an approximate ${}_1a_0$. The most common of these relationships, the Coale-Demeny and Keyfitz-Flieger formulas, rely on the general rule that as infant mortality falls, deaths become increasingly concentrated early in the interval, and ${}_1a_0$ also falls (Coale, Demeny, and Vaughan 1983; Keyfitz and Flieger 1971).[1] The Keyfitz-Flieger formula, for example, which is referenced in central demographic textbooks (e.g. Wachter 2014), is ${}_1a_0 = 0.07 + 1.71 {}_1M_0$, where ${}_1M_0$ is the infant mortality rate.

However, as noted by Andreev and Kingkade (2015), at the very low levels of infant mortality currently observed in the developed world, this is no longer the case, and in fact, ${}_1a_0$ has been rising in many countries since the 1970s. They attribute this to medical advances that reduce very early deaths due to congenital conditions and conditions of prematurity, and propose a new method for calculating ${}_1a_0$ based on a three-segment piecewise linear function that breaks mortality into extremely low, moderately low, and higher levels.

Andreev and Kingkade also note that the relationship between level of mortality and ${}_1a_0$ is rather weak at low levels of mortality. In this paper, we investigate the reason for this, and illustrate two competing effects on ${}_1a_0$: the overall level of infant mortality, and the share of births that are premature. Premature birth is difficult to predict and correlated with a number of factors, including maternal age, health and behavior, as well as socioeconomic status and race, and as such it may vary substantially, even within populations with relatively low infant mortality (Tucker and McGuire 2004; Purisch and Gyamfi-Bannerman 2017).

We illustrate the competing effect on the average age of infant death, ${}_1a_0$, through simulation, as well as an example with US data on two populations with very different levels of premature birth – infants born to non-Hispanic white mothers and infants born to non-Hispanic black mothers. Future work will focus on utilizing these observations in developing an alternative formula for ${}_1a_0$, which aims to capture not only the level of infant mortality but also the level of prematurity.

[1] This relationship can be explained with reference to Bourgeois-Pichat's theory (Bourgeois-Pichat 1951a; Bourgeois-Pichat 1951b), which assumes all infant deaths can be categorized as either endogenous or exogenous, with the latter tending to happen later in the first year. As overall mortality goes down, exogenous causes are increasingly eliminated, and the distribution of infant deaths shifts towards the endogenous causes, which are more likely to happen relatively soon after birth.

2 The relationship between ${}_1a_0$, infant mortality and prematurity

Although infant mortality relates to all deaths in the first year, the distribution of these deaths over time is far from uniform. For example, Figure 1 shows the distribution of infant deaths in the United States in 2012. The distributions have been plotted separately for pre- and full-term births, where preterm births are defined as those occurring before a gestational age of 37 full weeks from the last menstrual period.

Both figures show that irrespective of the prematurity of births, the largest share of infant deaths occurs in the first several days. The distribution is particularly skewed and concentrated for preterm births, with almost 20% of infant deaths occurring within the first 24 hours.

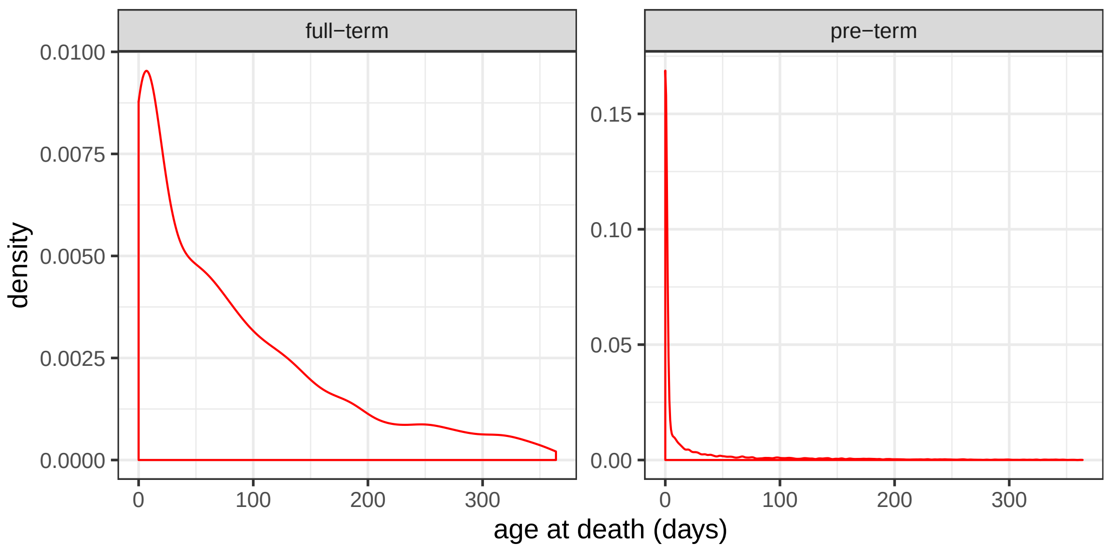

Figure 1: Density of infant deaths, United States, 2012 birth cohort. Data via National Center for Health Statistics National Vital Statistics System.

In terms of the relationship between overall infant mortality, prematurity and the distribution of death times, the following general observations can be made:

3. The distribution of the timing of infant deaths conditional on births being premature is also heavily left skewed, with an even larger density of deaths in the first few days (Figure 1 above).
4. The share of births that are premature tends to be higher in higher mortality conditions ("WHO Fact Sheet: Preterm Birth" 2018).

Why does the relationship between ${}_1a_0$ and infant mortality become unclear at lower mortality levels? Statements 1/2 and 3/4 above have competing effects on ${}_1a_0$. As overall mortality conditions improve, we expect that exogenous causes of death that occur later in the year decrease, and so ${}_1a_0$ will decrease. However, as overall mortality conditions improve, we also expect the share of births that are premature to decrease, and so ${}_1a_0$ will increase. These observations imply that trends in ${}_1a_0$ over time may go up or down; and that two populations with very different infant mortality rates may have the same or similar ${}_1a_0$

The competing effects are illustrated in this section; firstly by examining infants deaths in the United States by race, and secondly by simulating the effects on ${}_1a_0$ of changes in overall mortality risk and the prevalence of prematurity.

2.1 Example: US infant deaths by race

Racial disparities in U.S. infant mortality are longstanding and well-known. Although mortality has fallen for all racial and ethnic groups, babies born to non-Hispanic black mothers have consistently been dying at more than twice the rate of those born to non-Hispanic white mothers for over thirty years (Ely and Driscoll 2019; Hummer et al. 1999; Mathews, MacDorman, and Menacker 2013). Given this substantial inequality, it could be assumed that the average age of infant death, ${}_1a_0$, would be noticeably higher for the black population than the white population. However, as illustrated in this section, this is not the case. The disparity is not simply one of magnitude; patterns of preterm birth, low birth weight, age at death and cause of death all differ substantially between these two groups (Bediako, BeLue, and Hillemeier 2015; Ely and Driscoll 2019; MacDorman 2011; Riddell, Harper, and Kaufman 2017), and the interaction of these factors causes surprising patterns in ${}_1a_0$.

2.1.1 Data

Data on live births and deaths in the first year of life for infants born to non-Hispanic black (NHB) and non-Hispanic white (NHW) mothers comes from the NBER collection of U.S. Birth Cohort Linked Birth and Infant Death Data of the National Center for Health Statistics' National Vital Statistics System, years 2008-2012.[2] Cohort prematurity rates were calculated as the number of premature births divided by the total number of live births, and infant mortality rates were calculated as the number of deaths divided by the number of live births. Mortality was calculated according to race, birth cohort, premature status, age at death and cause of death. In line with WHO classification, premature status was grouped into 4 categories based on the last menstrual period (LMP) measure of gestational age: extremely preterm (born at $<$ 28 full weeks of gestation), very preterm (28 to $<$32 full weeks of gestation), later preterm (32 to $<$37 full weeks of gestation) and full-term (37 full weeks of gestation or more) ("WHO Fact Sheet: Preterm Birth" 2018). Age at death was split into first-week ($<$7 days old at death), neonatal ($<$28 days; includes first-week deaths), and post-neonatal (28 days or older at death). An average of 3021 births (0.1% of the total) and 147 deaths (0.8% of the total) were excluded each year due to missing gestational age

[2] National Center for Health Statistics. User Guide to the 2012 Cohort Linked Birth/Infant Death Public Use File. Hyattsville, MD: Centers for Disease Control and Prevention, US Department of Health and Human Services; 2018.

data. A replication data set including the raw data and statistical code is publicly available on GitHub (https://github.com/MJAlexander/a0_effects).

2.1.2 Mortality, prematurity and average age at death

Infant mortality rates for NHB and NHW populations in 2012, stratified by age at death and prematurity of birth, are shown below (Table 1).[3]

Large differences are observed in mortality rates by race that are not stratified by prematurity. Overall mortality is 2.21 times higher for infants born to NHB mothers than for those born to NHW mothers, at 11.00 per thousand versus 4.98 per thousand. When stratifying by age at death, a gradient is evident; racial ratios for mortality are highest in the first week (2.29), somewhat lower for neonatal mortality (2.24), and lowest, though still above 2, for post-neonatal mortality (2.15).

Among premature infants, mortality rates by gestational age are similar for those born to non-Hispanic black and non-Hispanic white mothers. For the extremely preterm, the ratio of the NHB mortality rate to the NHW mortality rate is 0.99. For the very preterm, the ratio is 1.00, and for later preterm infants, it is 1.04. For full-term births, however, NHB mortality is 1.70 times higher than NHW mortality.

		by age at death			by gestational age at birth			
	total infant mortality	first week	neonatal	post-neonatal	extremely preterm	very preterm	later preterm	full term
Black	11.00	5.94	7.30	3.70	363.18	38.86	9.01	3.54
White	4.98	2.60	3.26	1.72	365.32	38.90	8.65	2.09
Ratio	2.21	2.29	2.24	2.15	0.99	1.00	1.04	1.70

Table 1: Infant mortality rates for children of non-Hispanic black and non-Hispanic white mothers and the ratio between them

This discrepancy between patterns of mortality by gestational age and patterns by age at death is compositional; it is explained by a large difference in the distribution of births by gestational age, shown in Table 2. Fewer NHB infants are born at full term, and among those born early, NHB infants are more likely to be born extremely or very preterm. Because mortality risk drops rapidly with gestational age, this means that a larger share of NHB infants than NHW infants are at high risk of dying.

This compositional difference, in turn, leads to an instance of Simpson's paradox – wherein a trend observed in aggregate reverses when data are decomposed into subgroups. In aggregate, infants born to NHB mothers have a slightly lower average age at death, 40.87 days versus 45.29 days for infants of NHW mothers. This is somewhat surprising - as mentioned above, a longer ${}_1a_0$ is

[3] Because there is no large time trend in racial difference – that is, infant mortality is declining in a similar way for both racial groups – figures are given for the most recent birth cohort that data are available (2012). Figures for four previous cohorts may be found in Appendix A.

	Births (rates per thousand)			
	extremely preterm	very preterm	later preterm	full-term
Black	16.79	20.32	128.15	834.75
White	5.36	10.14	87.42	897.09
ratio	3.13	2.00	1.47	0.93

Table 2: Rates of preterm and full-term birth per 1000 live births, by race

generally associated with higher overall mortality and more preventable mortality, as deaths later in infancy are more likely to be caused by external factors and treatable diseases (Andreev and Kingkade 2015). However, within each subgroup by gestational age at birth, the average age at death is higher for NHB infants, indicating a greater proportion of deaths later in infancy. The difference is marked for all levels of prematurity; among full-term births, NHB infants' ${}_1a_0$ is only 3% longer than NHW infants'. (Table 3)

	mean age at death (days)				
	aggregate	extremely preterm	very preterm	later preterm	full term
Black	40.87	14.21	41.73	68.26	84.97
White	45.29	9.93	31.53	51.05	82.85
ratio	0.90	1.43	1.32	1.34	1.03

Table 3: Mean age at death (days)

2.2 Illustrating the competing effects through simulation

In order to further investigate the competing effects on ${}_1a_0$, we set up a simulation exercise that allows us to change the overall level of mortality and prematurity independently, and then assess the consequent change in ${}_1a_0$. Specifically, we carry out two simulations:

- Scenario 1: Vary the risk of infant mortality of a population with similar infant mortality to the full-term birth population in the US.
- Scenario 2: Vary the share of premature births in a population which has similar hazard profiles of pre- and full-term births in the US.

2.2.1 Simulation set-up

In order to be able to simulate plausible times of infant deaths, we need a suitable expression for the distribution of infant deaths. As illustrated in Figure 1, the distribution of infant deaths at the low levels of mortality we are interested in is highly skewed, with a large proportion of deaths occurring in the first week. The shape of these distributions is not readily captured by any classic parametric distributions. However, we found that the shape of infant death distributions was well-captured by a piecewise constant hazard (PCH) model, with time intervals partitioned at days 1-7, 14, 28, 60, 90, 180, and 365. The PCH model assumes constant exponential hazards within each of these time intervals, which allows the model to be estimated using Poisson regression. All estimation and calculations were carried out in R. Details on the statistical model and simulation can be found in Appendix B.

PCH models were fit separately to all full-term births/deaths and all preterm births/deaths using the 2012 US births and deaths data. Once estimates of hazards were obtained, the following process was used to simulate survival times and calculate ${}_1a_0$. For scenario 1:

1. Iterate through the risk factors [0.1,3] in increments of 0.1
2. Multiply the baseline hazards estimated from the full-term births model by the risk factor in 1.
3. Simulate death times from 10,000 births based on hazards in 2. (simulation details in Appendix B)
4. Repeat 1,000 times.
5. Save results and move to the next risk factor.
6. Calculate the mean, 2.5th and 97.5th quantiles of the average age of death for each risk factor.

For scenario 2:

1. Iterate through the shares of premature births between [0,0.5] in increments of 0.1. Call the current share s.
2. Simulate death times from $10,000 * s$ births based on hazards estimated from preterm births model.
3. Simulate death times from $10,000 * (1 - s)$ births based on hazards estimated from full-term births model.
4. Repeat 1,000 times.
5. Save results and move to the next risk share.
6. Calculate the mean, 2.5th and 97.5th quantiles of the average age of death for each prematurity share.

2.2.2 Results

The results of these simulations are shown in Figures 2 and 3. The results clearly show that changes in overall infant mortality and changes in prematurity act in opposite directions. For increases in the risk profile of mortality, the average age of infant death steadily increases until a risk factor of around 2, when the estimated ${}_1a_0$ plateaus. This plateau is a consequence of the shape of the death distribution, with increases in later-infant mortality being balanced out by the increases in high hazards in the earlier-infant mortality. For increases in the share of premature deaths (3), the average age of infant death steadily declines. It is worth noting that the magnitude of changes in ${}_1a_0$ in response to changes in prematurity are larger than those based on increasing the overall risk of mortality. This is broadly consistent with the pattern observed in the U.S. data; faced with both higher rates of prematurity and a higher overall risk of mortality, NHB have a shorter ${}_1a_0$ than NHW infants.

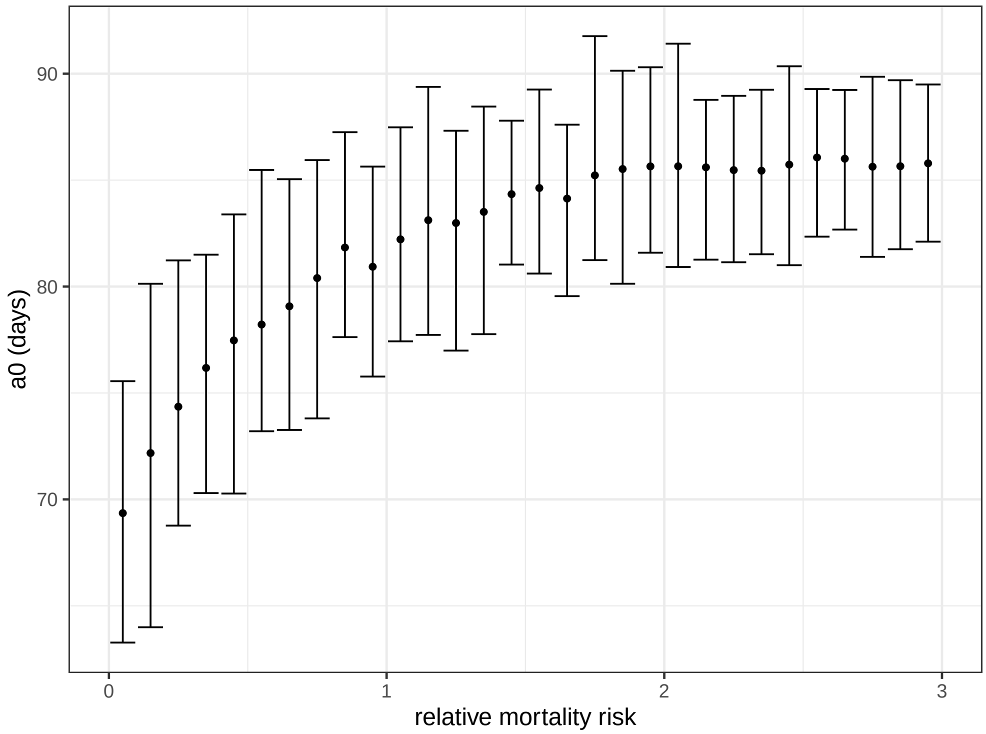

Figure 2: Estimated ${}_1a_0$ and 95% confidence intervals for different risk factors. Results from 1,000 simulations of 10,000 births with baseline hazards fitted to the NHB population in 2012.

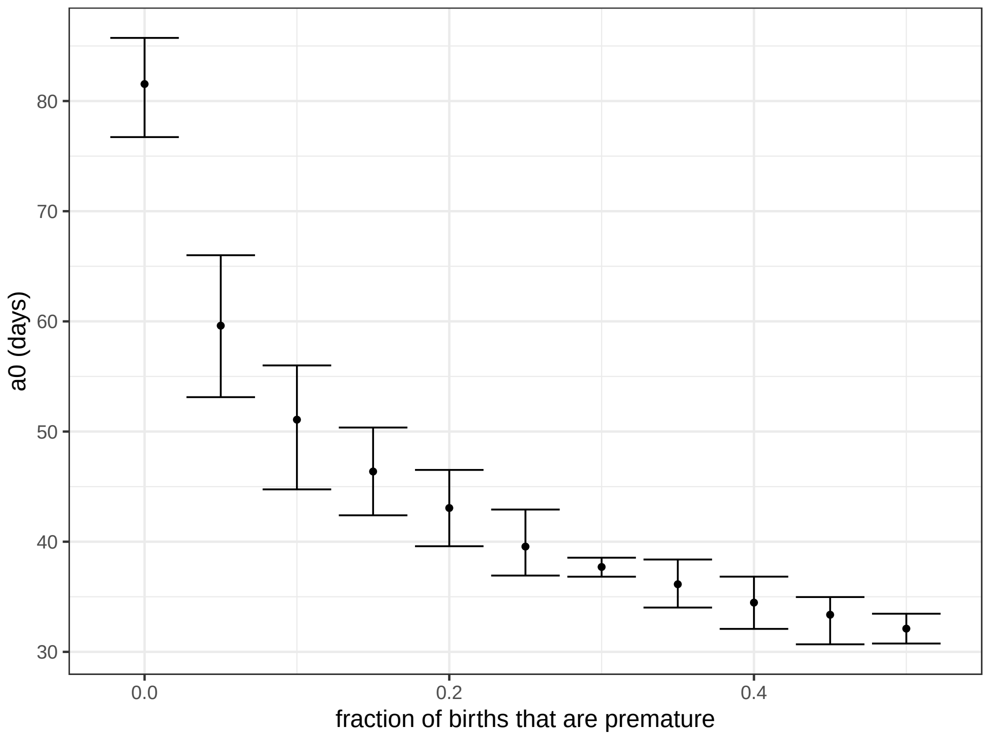

Figure 3: Estimated ${}_1a_0$ and 95% confidence intervals for different shares of premature births. Results from 1,000 simulations of 10,000 births with baseline hazards for pre- and full-term births fitted to the US population in 2012.

3 Towards a new formula for ${}_1a_0$

Future work will focus on developing alternative formulas for estimating ${}_1a_0$ that incorporate both mortality and prematurity effects. Because rates of prematurity are not readily available for all countries, we will investigate the possibility of including the neonatal mortality rate and/or the ratio of neonatal to other infant deaths in models, as a proxy for prematurity. Model building and testing will focus on using the detailed US data available, and other data sources will be investigated to generalize findings to other populations.

4 Summary and future work

Past research has noted both increasing variation, as well as a general rise, in ${}_1a_0$, the mean length of life for infants who die (Andreev and Kingkade 2015). In this paper we have shown that, at low levels of mortality, this shift can be explained by the fact that a reduction in the share of preterm births and an overall decline in mortality have competing effects on ${}_1a_0$, with the former increasing it and the latter decreasing it. This is demonstrated both empirically using U.S. infant mortality data, and through a simulation exercise that varies the overall level of mortality and prematurity independently.

Although minor variations in ${}_1a_0$ are unlikely to significantly affect important lifetable measures such as life expectancy (e_0), incorporating prematurity rates into its calculation could help restore its relationship to overall mortality conditions and improve its utility as an indicator of population well-being. For this reason, we propose the development of alternative estimation formulas in the future, incorporating empirical data on prematurity and/or neonatal mortality.

A Extra tables

Tables below show mortality rates and prematurity rates for the US for the years 2008–2012.

	2008	2009	2010	2011	2012
	infant mortality rate				
Black	12.34	12.03	11.32	11.07	11.00
White	5.40	5.19	5.06	4.98	4.98
ratio	2.29	2.32	2.24	2.22	2.21
	first-week mortality rate				
Black	6.48	6.33	5.85	6.03	5.94
White	2.71	2.59	2.56	2.57	2.60
ratio	2.39	2.44	2.28	2.35	2.29
	neonatal mortality rate				
Black	8.07	7.93	7.28	7.43	7.30
White	3.44	3.32	3.29	3.25	3.26
ratio	2.35	2.39	2.21	2.29	2.24
	post-neonatal mortality rate				
Black	4.27	4.09	4.04	3.64	3.70
White	1.96	1.87	1.77	1.74	1.72
ratio	2.18	2.19	2.29	2.10	2.15

Table 4: Mortality rates by age at death, race, and year

	2008	2009	2010	2011	2012
	extremely preterm				
Black	398.48	384.46	366.24	362.18	363.18
White	387.33	376.18	369.73	369.00	365.32
ratio	1.03	1.02	0.99	0.98	0.99
	very preterm				
Black	46.62	45.77	44.10	40.36	38.86
White	40.26	41.40	37.39	41.58	38.90
ratio	1.16	1.11	1.18	0.97	1.00
	later preterm				
Black	10.46	9.92	9.67	9.88	9.01
White	8.84	8.18	8.30	8.26	8.65
ratio	1.18	1.21	1.16	1.20	1.04
	full term				
Black	3.72	3.64	3.53	3.36	3.54
White	2.29	2.22	2.13	2.11	2.09
ratio	1.63	1.64	1.66	1.60	1.70

Table 5: Mortality rates by premature status, race, and year

	2008	2009	2010	2011	2012
	extremely preterm				
Black	17.19	17.43	16.89	16.99	16.79
White	5.41	5.36	5.44	5.24	5.36
ratio	3.18	3.25	3.10	3.25	3.13
	very preterm				
Black	21.24	21.25	20.99	20.58	20.32
White	10.58	10.38	10.36	10.21	10.14
ratio	2.01	2.05	2.03	2.02	2.00
	later preterm				
Black	137.00	135.98	133.30	130.17	128.15
White	95.39	93.43	91.89	89.56	87.42
ratio	1.44	1.46	1.45	1.45	1.47
	full term				
Black	824.57	825.34	828.82	832.26	834.75
White	888.62	890.82	892.30	895.00	897.09
ratio	0.93	0.93	0.93	0.93	0.93

Table 6: Rates of preterm and full-term birth per 1000 live births by race and year

B Details on simulation

B.1 Goals

The goal of the simulation exercise is to see how ${}_1a_0$ varies with changes in 1) the overall mortality risk and 2) the share of births that are premature. In order to do this, we need to first obtain a set of hazards for full-term births a preterm births. We then need to obtain a set of simulated death times based on these hazards under the conditions of 1) and 2).

B.2 Estimating hazards

In order to simulate infant death times under varying mortality risk and prematurity conditions, an expression for the hazard of dying at a particular time t is required. However, the shape of the distribution of infant death times is non-parametric and thus not well represented by standard parametric distributions like the Exponential or Weibull.

As such, in order to obtain estimates for the hazards over time, we used piecewise constant hazard (PCH) models. We assume exponential hazards within each piece, and split the first year at days 1-7, 14, 28, 60, 90, 180, and 365. It can be shown that piece-wise exponential proportional hazards model is equivalent to a Poisson log-linear model where the death indicator is the response and the log of exposure time enters as an offset. We fit these models in `R` using the `glm` function separately for all full-term births and all preterm births in the 2012 US dataset. The resulting regression estimates give the log-hazards in each of the separate intervals, which can then be used for simulation.

B.3 Piecewise exponential survival function

Define the set of time intervals to be $0 = \tau_0 < \tau_1 < \dots \tau_m < \tau_{m+1}$. The piecewise constant hazard function is

$$h(t) = h_0 \sum_{l=0}^{m} g_l I_l(t)$$

where h_0 is the baseline hazard, g_l are the relative hazards and $I(t)$ is an indicator function which is equal to 1 if $\tau_l \leq t < \tau_{l+1}$ and 0 otherwise.

The survival function is

$$S(t) = \exp\left(-H(t)\right) = \exp\left(-h_0 \sum_{l=0}^{m} g_l \int_0^t I_l(s) ds\right)$$

The piecewise survival function for interval if $\tau_i \leq t < \tau_{i+1}$ is

$$S_i(t) = \exp\left(-h_0\left(\sum_{l=0}^{i-1} g_l(\tau_{l+1} - \tau_l) + g_i(t - t_i)\right)\right)$$

Solving for t implies

$$t = \tau_i - \frac{\log(S_i(t))}{h_0 g_i} - \frac{1}{g_i}\sum_{l=0}^{i-1} g_l(\tau_{l+1} - \tau_l) \tag{1}$$

As $\tau_i \leq t < \tau_{i+1}$ this gives us two conditions:

$$\log(S_i(t)) \leq -h_0 \sum_{l=0}^{i-1} g_l(\tau_{l+1} - \tau_l) \tag{2}$$

$$\log(S_i(t)) > \sum_{l=0}^{i} g_l(\tau_{l+1} - \tau_l) \tag{3}$$

B.4 Simulation framework

We have an analytical form for the survival function S, which means there is also an analytical form for the cumulative distribution function $F = 1 - S$. We used inverse transform sampling and equations 1-3 above to generate samples from the death distributions specified by a set of piecewise hazards estimated using Poisson regression. To reiterate, the two simulation scenarios were:

- Scenario 1: Vary the risk of infant mortality of a population with similar infant mortality to the full-term birth population in the US. Risk multipliers simulated were between [0.1,3] in increments of 0.1.
- Scenario 2: Vary the share of premature births in a population which has similar hazard profiles of pre- and full-term births in the US. Shares considered were between [0, 0.5] in increments of 0.1.

The following simulation process was used:

1. Set the time intervals, the baseline hazard, and all relative hazards. For Scenario 1, the hazards come from the full-term birth estimation. For Scenario 2, both the hazards from full- and preterm births are used. For every iteration in each scenario (i.e. each risk multiplier and each share s)

2. Set the size of the sub-population. For Scenario 1 this was 10,000. For Scenario 2 this was $10,000 * s$ preterm births and $10,000 * (1 - s)$ full-term births.

3. Draw a uniformly (0,1) distributed random variable S = 1 - F. (For Scenario 2, two random variables are drawn.)

4. Determine the right interval using the conditions in Equations 2 and 3.
5. Compute the random time t using Equation 1.
6. Combine the computed random times to one file.
7. Repeat this process 1,000 times.

Once the set of 1,000 simulation times are obtained, the mean age of death ${}_1a_0$ was calculated for each simulation. The mean, 2.5th and 97.5th quantiles of ${}_1a_0$ were then calculated.

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