

# WAITING PERIOD FROM DIAGNOSIS FOR MORTGAGE INSURANCE ISSUED TO CANCER SURVIVORS

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# Waiting period from diagnosis for mortgage insurance issued to cancer survivors

by Antoine Soetewey

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*“You can’t connect the dots looking forward; you can only connect them looking backwards. So you have to trust that the dots will somehow connect in your future.”*  
(STEVE JOBS)



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*Antoine Soetewey*  
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# Bibliographic notes

## Journal publications

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- Truong, H. C., et al. Childhood Bacterial Meningitis Surveillance in Southern Vietnam: Trends and Vaccination Implications from 2012 to 2021. *Open Forum Infectious Diseases*, 10(7):ofad229. Oxford University Press, 2023. <https://doi.org/10.1093/ofid/ofad229>
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- Workshop: “How can we predict the evolution of Covid-19 in Belgium?” (2020). UCLouvain, Belgium [slides]
- Scientific exchange day: “Cure models: estimating the recovery time to improve access to insurance” (2020). French National Cancer Institute (INCa), France [slides]
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### **Miscellaneous contributions**

- How can we predict the evolution of COVID 19 in Belgium? [LIDAM communication]
- The right to be forgotten - Are we all equal when it comes to insurance? [LIDAM news]

# List of acronyms

AERAS	S'Assurer et Emprunter avec un Risque Aggravé de Santé
BCR	Belgian Cancer Registry
BH	Baseline hazard
BRCA1/BRCA2	Breast cancer 1/Breast cancer 2
CI	Confidence interval
EPV	Expected present value
EU	European Union
FRANCIM	French network of cancer registries
FSMA	Financial Services and Markets Authority
HIV	Human Immunodeficiency Virus
ICD-10	International Classification of Diseases, Tenth Revision
KCE	Belgian Health Care Knowledge Centre
KM	Kaplan-Meier
LNPH	Linear and non-proportional hazard
LPH	Linear and proportional hazard
MSM	Multi-state model
NBB	National Bank of Belgium
NISS/INSZ	Numéro d'Identification à la Sécurité Sociale/IdentificatieNummer Sociale Zekerheid
NLNPH	Non-linear and non-proportional hazard
NLPH	Non-linear and proportional hazard
RTBF	Right to be forgotten
SE	Standard error
TTC	Time-to-cure
UK	United Kingdom
YLL	Years of life lost



# Introduction

# | 1

Property loans are often accompanied with mortgage insurance that pays the balance of the loan if the mortgagor dies. Coverage is usually awarded in the form of term insurance with decreasing sum insured, with the amount of death benefit diminishing as the debt decreases. This is common practice in Belgium, with about 170,000 new mortgage loans per year, mainly contracted by young adults acquiring their first family house (statistics from the Belgian Central Credit Register indicate that 36% of new mortgage loans in 2017 were contracted by borrowers younger than 35 and about 68% were granted to borrowers younger than 45).

Based on answers to a health questionnaire, insurers evaluate applicant's health status. As with any other term life insurance product, applicants with poor health conditions may be denied insurance or charged increased amounts of premium compared to standard conditions. In extreme cases, this may prevent them from accessing property (in case of a house loan) or develop their business project (in case of a professional loan). Filling such health questionnaires may create frustration for patients having survived cancer and which was diagnosed many years ago. Having repeatedly to answer questions related to this disease has psychological consequences and being charged higher premiums or denied coverage generates a feeling of discrimination (Massart, 2018). This is often felt as a double penalty by cancer survivors. For this reason, several EU countries passed laws to ease access to mortgage insurance for long-term disease survivors. This materializes into the "right to be forgotten" adopted in several EU member states, granting access to insurance after a waiting period of at most 10 years starting at the end of the successful therapeutic protocol.

Combining concepts from biostatistics and actuarial science, there are two topics that dominate the thesis. The first one is related to the right to be forgotten in insurance. In particular, the aim is to develop a method to adequately estimate the threshold after which cancer patients can be considered as cured. For some types of cancer, survivors actually have a chance of survival comparable to that of the general population only after a few years after diagnosis, or pose a moderately increased risk and could therefore be covered in the event of death. This involves measuring and quantifying the potential excess mortality so that the premiums claimed reflect the risk in terms of financial services. The second dominant topic of the thesis is to find a proper way to adapt the actuarial pricing of life insurance products to each category of risk, disease, person, etc.

The first problem has been studied in different contexts in the literature. Our contribution is that the right to be forgotten starting from the date of diagnosis, instead of the date of the end of the therapeutic treatment, is presented. The second problem is tackled by concentrating on pricing insurance covers on a market where such a right has been implemented. To better clarify the two problems in the focus of the thesis, an introduction of the right to be forgotten in Belgium and other European countries is needed.

The first initiative dates back to 2007, when France launched the AERAS Convention (AERAS is the acronym for "*s'Assurer et Emprunter avec un Risque Aggravé de*

*Santé*” in French, which could be translated as “insuring and borrowing under poor health conditions”). This agreement, signed by the public authorities, banking and insurance sectors, and patients’ and consumers’ associations purposed to allow people cured from cancer or suffering certain chronic diseases to access insurance in case of death or disability, as well as to guaranteed income insurance. Faced with a similar situation, this lead Belgian authorities to create the *Bureau du suivi de la tarification assurance solde restant dû* ([www.bureaudusuivi.be](http://www.bureaudusuivi.be)) – *Opvolgingsbureau voor de tarifiering schuldsaldoverzekerings* ([www.opvolgingsbureau.be](http://www.opvolgingsbureau.be)) (which could be translated literally as the “Outstanding balance insurance pricing monitoring office”) in 2014, in application of the law on insurance. This body reviews health questionnaires used by insurance companies selling mortgage insurance in Belgium and checks whether the proposed premium surcharges or cover denials are justified for impaired lives. In 2017, only 16% of the 454 cases submitted by insurance applicants to the Bureau resulted in improved policy conditions, as it can be read from the annual report published by the Bureau (2018).

Considering long-term cancer survivors, France established in 2016 a “*droit à l’oubli*” (translated literally as “right to be forgotten” (RTBF) in the remainder of this text), that is, the right for an insurance applicant not to declare a previous cancer after a period of 10 years starting at the end of the therapeutic protocol. This 10-year waiting period is reduced to 5 years if the applicant suffered cancer before the age of 18. These periods of 10 and 5 years start from the date of the end of the therapeutic treatment, in absence of relapse within this period. The 10-year length of the waiting period is further shortened for several types of cancer (and other non cancer related pathologies), as detailed in the AERAS Convention (see [www.aeras-infos.fr/cms/sites/aeras/accueil.html](http://www.aeras-infos.fr/cms/sites/aeras/accueil.html)) with reduced duration after which survivors have access to the right to be forgotten. After this period, insurance companies cannot take the pathology into account in risk assessment and cannot refuse the insurance, nor impose a premium surcharge because of the pathology.

It was only in 2019 that the RTBF entered the insurance law in Belgium. Based to a large extent on the reference tables published in the AERAS Convention, a Royal Decree dated May 26, 2019 lists certain types of cancer for which, depending upon entry criteria (such as cancer stage or age), the standard waiting period of 10 years from the end of active treatment is reduced. The RTBF has recently been adapted in Belgium, again following similar changes in France. The standard waiting period opening the RTBF has been shortened from 10 years to 8 years (as from November 2022) and to 5 years (as from January 2025), or to 5 years (as from November 2022) if the insurance applicant was diagnosed with the considered cancer before the age of 21.

Every two years, the Belgian Healthcare Knowledge Centre (KCE) evaluates the reference grid in the light of medical progress and available scientific data on the pathologies it covers. The KCE then proposes an adaptation of the reference grid, which it communicates to the *Bureau du suivi de la tarification*. The latter forwards the proposal, together with its opinion, to the Minister for Insurance and the Minister for Social Affairs. The reference grid may then be modified if necessary.

The RTBF has now also been installed through an agreement (convention) in Luxembourg, and through a legal framework in The Netherlands, Portugal, Romania, and more recently in Italy, Spain and Cyprus. Established in 2020 in Luxembourg, the agreement applies exclusively to the outstanding balance of an insurance policy for the acquisition of a principal residence or business premises, up to a maximum amount of 1,000,000€. It does not apply to the acquisition of a second home or to rental investments. Established in 2021 in The Netherlands, the law applies to life insurance for applicants under the age of 71. The decree establishes that it is no

longer allowed to ask whether someone has already had cancer when, in the opinion of the health care provider who treated the applicant, there has been a complete remission and no recurrence has been diagnosed for an uninterrupted period of 10 years, starting from the moment when complete remission was established. The decree also states that if the candidate is aged under 21 at the time of cancer diagnosis, the period is reduced to 5 years. Established in 2022 in Portugal, the law gives people who have overcome serious illnesses such as cancer, HIV and diabetes the right to be forgotten when taking out a mortgage, consumer credit or insurance (whether compulsory or optional) linked to these loans, provided that 10 years have elapsed without interruption since the end of the therapeutic protocol, and 5 years have elapsed since the end of the therapeutic protocol in the case the illness was diagnosed before the age of 21. Established in 2022 in Romania, the right to be forgotten applies to adults 5 years after the end of treatment in case the cancer diagnosis occurred before the age of 18 years, and after 7 years following the end of their treatment and without any evidence of relapse or recurrence in case the cancer diagnosis occurred after the age of 18 years. Established in 2023 in Spain, the right to be forgotten in the field of insurance and banking products for cancer patients becomes effective once 5 years have elapsed since the end of treatment without relapse. Also in 2023, Italy passed a law allowing cancer survivors not to share information about their previous condition with financial institutions, or adoption authorities, provided that 10 years had passed since the successful end of their treatment, or 5 years in the case of cancer diagnosed before the age of 21. In November 2023, Cyprus followed the trend by approving a law that states that no insurance company will be able to reject an application made by a cancer survivor if 10 years or more have passed since the completion of their treatment, or 5 years in the case of cancer diagnosed before the age of 21.

Besides the implementation of the right to be forgotten in these countries, it is being debated and advocated for at the European level to expand to the other EU countries as well. There are some ongoing discussions between Insurance Europe, the European Commission and the European Parliament on a possible EU-wide RTBF for cancer survivors. See for instance Scocca and Meunier (2020, 2022), as well as to “Survivorship challenge 3.4: Lack of knowledge of the stigma associated with cancer” listed in Lawler et al. (2021) purposing to take advantage of the existing legal framework in four EU member countries (France, Belgium, Luxembourg and the Netherlands) to investigate a pan-European legal framework on access to financial services for cancer survivors.

## 1.1 Summary of the chapters

The manuscript is divided into several chapters. The main findings and contributions of each of the chapters are summarized.

**Chapter 2: Waiting period from diagnosis for mortgage insurance issued to cancer survivors.** This chapter, identical to our first paper (Soetewey et al., 2021), focuses on the application of several tools from biostatistics to assess the length of the waiting period opening the right to be forgotten. Although the establishment of such a right to be forgotten in several European countries and possibly at the European level is clearly an improvement for cancer survivors, it is shown, using data from the Belgian Cancer Registry, that there is most probably room for further reducing the waiting period for some cancer types. In particular, the chapter aims to show that for some types of cancer (with melanoma and thyroid as examples), survivors actually have a survival comparable to that of the general population, that is, excess

mortality is negligible. It is also demonstrated that patients having survived long enough to some types of cancer (still with melanoma and thyroid as examples) can access life insurance market at standard insurance rates, contrarily to the common belief within the actuarial community. Moreover, there remains some ambiguity about what is considered as treatment and thus what marks the end of the therapeutic protocol. Therefore, a waiting period starting at diagnosis rather than at the end of the therapeutic treatment protocol is promoted in order to avoid disputes in case of death. Results appear to be particularly encouraging as they suggest a considerable shortening of the 10-year waiting period for some types of cancer.

**Chapter 3: Semi-Markov modeling for cancer insurance.** This chapter, identical to our second paper (Soetewey et al., 2022), focuses on insurance covers on a market where a right to be forgotten has been implemented. More precisely, the products considered here are specifically related to the waiting period opening the right to be forgotten, with temporary covers restricted to that period to fill the gap in coverage on a market where such a right has been implemented. First, stand-alone products are studied, including cancer insurance with lump sum payment at diagnosis, or temporary life annuity starting at diagnosis. In the latter case, periodic payments may correspond to insurance premiums of another product, or even to loan reimbursement. Then, riders included in a package are discussed. Term insurance with accelerated death benefit paid as a lump sum at diagnosis or as a temporary life annuity starting at diagnosis are considered. Finally, products granting access to some specific insurance cover (such as mortgage insurance) during the waiting period opening the right to be forgotten are discussed. This is especially important at young age, to guarantee access to property and home ownership (in case of house loan) and to entrepreneurship (in case of professional loan) to cancer patients whose health status has improved but who cannot benefit from the right to be forgotten because the waiting period is not exhausted. The 3-state (healthy–ill–dead) Semi-Markov hierarchical model developed in Denuit et al. (2019) for long-term care insurance is adopted here for actuarial calculations. Semi-Markov transition intensities are estimated from cancer cases recorded by the Belgian Cancer Registry. Our proposals are illustrated through three cancers with clear differences in terms of incidence, survival after diagnosis, and waiting periods defined by Royal Decree: (i) melanoma, (ii) thyroid and (iii) female breast cancers. The obtained results suggest that a new offer could develop, targeting the particular needs of cancer patients.

**Chapter 4: Health indices for disease incidence risk and duration in the Semi-Markov setting.** This chapter, identical to our third paper (?), focuses on illustrating how common health indices (with disease incidence risk and years of life lost as examples) can be estimated based on a Semi-Markov three-state illness-death model using cancer registry data. The main advantage of computing these quantities in a Semi-Markov context is that it allows to take into account the number of years a patient survived after the diagnosis. To the best of our knowledge, most studies refer to the number of years of life lost at the time of diagnosis, without taking the time survived since diagnosis into consideration. This is a major difference, given that time survived since diagnosis is known to have an influence on survival for cancer patients. Based on 161,007 melanoma, thyroid and female breast cancer cases recorded by the Belgian Cancer Registry, it appears that the probabilities of being diagnosed with cancer over the next 20 years for a healthy individual remain rather low for melanoma and thyroid cancers for both sexes, but considerably increases with age for female breast cancer. It also turns out that the number of years of life lost due to cancer (i) is lowest for women with thyroid cancer and highest for women with breast cancer, (ii) decreases with age at diagnosis for both sexes and all three cancers



under consideration, and (iii) is larger for men than for women for both melanoma and thyroid cancers. xxx adapt when paper is published.

**Chapter 5: Right to be forgotten for mortgage insurance issued to cancer survivors: Critical assessment and new proposal.** This chapter, identical to our fourth paper (?), is a follow-up of the first chapter. In the first chapter, it has been proposed to determine the waiting period opening the right to be forgotten as the time after diagnosis needed for the premium to revert back to some acceptable level, expressed by means of regulatory life tables. However, this approach requires data up to 30 years after diagnosis (10 years of standard right to be forgotten plus the typical duration of the loan), or extrapolating the results up to that time horizon. In this chapter, it is shown that when survival statistics are only available over a shorter duration, it turns out that the length of the resulting waiting period opening the right to be forgotten may strongly depend on the extrapolation method. This problem, not arising from the method proposed in the first chapter but coming from the limited follow-up period for patients in some cancer registries (including the Belgian one), is not acceptable in the context of the right to be forgotten. This is why an alternative method is proposed here, based on a constraint imposed to the premium. This constraint is then transposed into a target on the conditional observed survival probabilities. The length of the waiting period opening the right to be forgotten can then be derived from the comparison of the conditional one-year survival probabilities of cancer patients with the corresponding probabilities at general population level. The main advantage is that the time from which the right to be forgotten can be exercised can be estimated from the available data only, without the need to extrapolate mortality rates beyond 10 years. For the sake of robustness, results obtained with the proposed approach are compared to results obtained with Kaplan-Meier estimate taken as a non-parametric reference. Furthermore, while cancer stage at diagnosis has not been taken into account in the first chapter, the impact of the stage of the tumor at diagnosis on waiting periods is investigated in this chapter. xxx adapt when paper is published.

**Chapter 6: Conclusion.** In the last chapter, contributions of this manuscript are briefly summarized. Furthermore, some problems that could not be solved during the thesis and which possibly already appeared in the previous chapters are listed. Some general questions that arose while working on the topics from the previous chapters, and whose solution might either have important practical or conceptual implications for both biostatisticians and actuaries, are then proposed.

## 1.2 Prerequisites

Before delving into the different chapters of this PhD thesis, it is recommended for the reader to possess a foundational understanding of several key concepts and methodologies. These prerequisites are carefully curated to ensure a seamless comprehension of the intricate topics covered in the thesis.

### 1.2.1 Survival analysis

Survival analysis is a branch of statistics aimed at analyzing the expected duration of time until one or more events happen. In survival analysis, we are interested in a certain event and we would like to analyze the time until the event happens (referred as survival time).

Although survival analysis is rooted in medical and public health applications (with the time to death), it is now used in many domains. For example, one may also be interested in the time until:

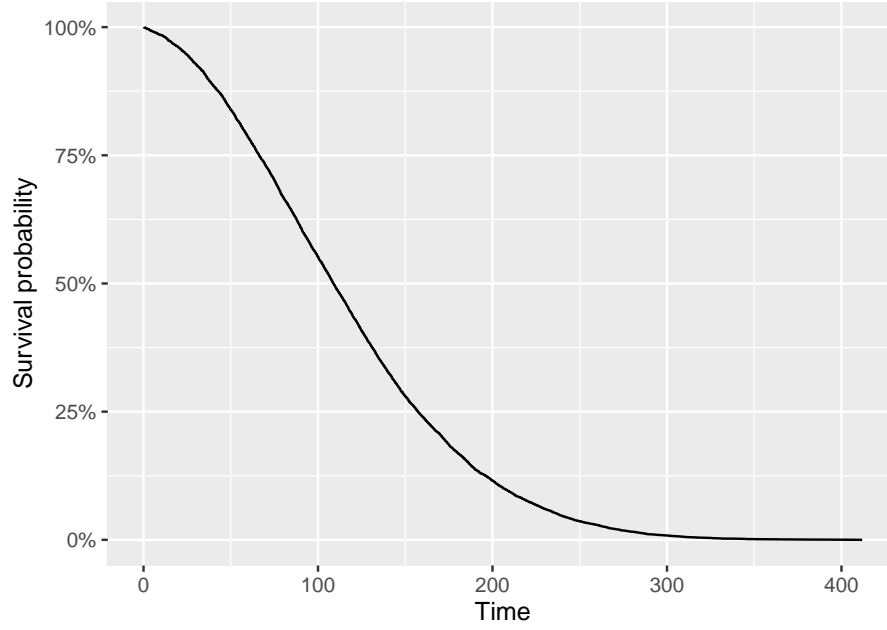
- an unemployed person finds a job,
- a citizen is being arrested again after having been released from jail,
- a woman becomes pregnant for the first time,
- a machine breaks down,
- a company goes bankrupt,
- a customer buys a new product or stops its current subscription,
- a letter is delivered,
- a taxi picks you up after having called the taxi company,
- an employee leaves the company,
- etc.

As it can be seen, the event of interest does not necessarily have to be the death or the occurrence of a disease, but in all situations we are interested in analyzing the time until a specific event occurs.

Survival data requires a special set of statistical methods for three main reasons:

1. Survival times are always positive. The time until an event of interest occurs cannot be less than 0.
2. Different measures are of interest depending on the research question and the context. For instance, one could be interested in knowing (i) the probability that a cancer patient survives longer than 5 years after diagnosis, (ii) the typical waiting time for a cab to arrive after having called the taxi company, or, (iii) out of 100 unemployed people, how many are expected to have a job again after 2 months of unemployment.
3. Censoring is almost always an issue. Most of the time, the event occurred before the end of the study and the survival time is known. However, sometimes, the event is not yet observed at the end of the study. Suppose that we study the time until death of patients with breast cancer. Luckily, some patients will not die before the end of the study. Other times, another event occurs before the event of interest which prevents it from ever happening (for example, a cancer patient may die from a car accident). It can also happen that a patient withdraws from the study or moves to another country; his or her survival time cannot be observed. Censoring can be seen, in some sense, as a type of missing data. For these reasons, many standard statistical tools are not appropriate for this type of data. Specific statistical methods are required to take into account the fact that the exact survival time for some patients is missing. It is known that they survived a certain amount of time (until the end of the study or until the time of withdrawal), but their exact survival time is unknown.

The two most common types of censoring are right-censoring and left-censoring. When the event is not yet observed at the end of the study (i.e., the survival duration is greater than the observed duration), this is referred as right-censoring. On the other hand, left-censoring occurs if a participant is entered into the study when the



**Figure 1.1: Example of a survival function  $S(t)$**

event of interest occurred prior to study entry but we do not know exactly when. The goal is of course to analyze all available data, including information about censored patients. The aim of survival analysis is thus to model and describe survival data in an appropriate way, taking the particularities of this type of data into account.

Without going too much into the details, we lay the foundation with the most common functions in survival analysis:

- the survival function,
- the cumulative hazard function, and
- the hazard function.

#### **1.2.1.1 Survival function**

Let  $T$  be a non-negative continuous random variable, representing the time until the event of interest. For each  $t$ :

$$S(t) = P(T > t) = 1 - F(t).$$

In the context of this thesis, it gives the probability that a randomly selected cancer patient will survive beyond time  $t$ . The survival function  $S(t)$  is a decreasing function taking values in  $[0, 1]$ , and equal to 1 at  $t = 0$  and 0 at  $t = \infty$ . Figure 1.1 shows an example of a survival function. The curve gives the proportion of individuals (or experimental units) who, as time goes on, have not experienced the event of interest. As time progresses, events occur, so the proportion who have not experienced the event decreases.

**1.2.1.2 Cumulative hazard function**

The cumulative hazard function, which corresponds to the total hazard experienced up to time  $t$ , is defined as:

$$\Lambda(t) = -\log S(t).$$

The cumulative hazard function is an increasing function taking value in  $[0, +\infty]$ , and  $S(t) = \exp(-\Lambda(t))$ .

**1.2.1.3 Hazard function**

The hazard function, or hazard rate, is defined as:

$$\begin{aligned}\lambda(t) &= \lim_{\Delta t \rightarrow 0} \frac{P(t \leq T < t + \Delta t | T \geq t)}{\Delta t} \\ &= \frac{d}{dt} \Lambda(t).\end{aligned}$$

The hazard function is a positive function (not necessarily increasing or decreasing), and it can have many different shapes (which is therefore a useful tool to summarize survival data). In the context of cancer patients,  $\lambda(t)$  measures the instantaneous risk of dying right after time  $t$  given the individual is alive at time  $t$ .

To link the hazard rate with the survival function; the survival curve represents the hazard rates. A steeper slope indicates a higher hazard rate because events happen more frequently, reducing the proportion of individuals who have not experienced the event at a faster rate. On the contrary, a gradual and flatter slope indicates a lower hazard rate because events occur less frequently, reducing the proportion of individuals who have not experienced the event at a slower rate. Note that, in contrast to the survival function which focuses on not having an event, the hazard function focuses on the event occurring.

**1.2.1.4 Estimation of the survival function and overall survival**

To estimate the survival function, an estimator which is able to deal with censoring is needed. The most common one is the nonparametric Kaplan-Meier (1958) estimator. The advantages of this estimator are that:

- it is simple and straightforward to use and interpret,
- it is a nonparametric estimator, so it constructs a survival curve from the data and no assumptions is made about the shape of the underlying distribution, and
- it gives a graphical representation of the survival function(s), useful for illustrative purposes.

Note that an important assumption for the estimation to hold is that censoring is independent of the occurrence of events. This is referred as non-informative censoring, and means that censored subjects have the same survival prospects as subjects who are not censored and who continue to be followed.

In order to estimate a survival function with the Kaplan-Meier estimator, the following two pieces of information are required:

1. the time until the event of interest or the time until the censoring, and
2. the event status (whether the event happened or not).

The survival curve resulting from the Kaplan-Meier estimator can be seen as a descriptive statistic for survival data. This estimator is used to estimate the overall survival function  $S(t)$ , without distinguishing according to causes of death (Belot et al., 2019). Several quantities can be obtained from this estimator, such as the median survival time (i.e., the time at which the survival function  $S(t)$  is 50%) or the probability that a cancer patient will survive, say, more than 1 or 5 years after diagnosis (known as the 1 and 5-year survival probability). Another important concept in survival analysis is the relative survival, which is detailed in the next section.

#### 1.2.1.5 Relative and net survival

Information on cause of death due to cancer is often unavailable or unreliable and not all deaths can be easily classified as deaths due to the cancer of interest or due to another cause (Percy et al., 1981). Relative survival, which does not require information on the cause of death, provides a measure of the excess mortality experienced by cancer patients by comparing the mortality in the cancer population with the mortality in the general population. This led the relative survival to become the standard measure of patient survival for population-based cancer registries, as shown by its prominence around the world in studies related to cancer survival (Ries et al., 2002; Coleman et al., 1999; Berrino et al., 1999; Perme et al., 2012, 2016; Pavlič and Pohar Perme, 2019).

The hazard rate at time  $t$  since diagnosis for cancer patients with covariate vector  $Z$ , is decomposed into two additive components: the population hazard based on available patient's characteristics  $Z = z$ , denoted as  $\lambda_P(t|z)$ , and the excess hazard specific for the disease of interest, denoted as  $\lambda_E(t|z)$ . The relative survival function  $r(t|z)$  corresponds to the ratio of the survival function of the studied group  $S(t|z)$  to the survival function of a comparable group (i.e., with the same characteristics) from the general population  $S_P(t|z)$  (Dickman et al., 2004):

$$\frac{S(t|z)}{S_P(t|z)}.$$

The population hazard  $\lambda_P(t|z)$  is usually estimated on the basis of external data such as population life tables, which are usually stratified according to the main factors affecting patient survival such as age, gender, and calendar year.

Net survival is a measure of patient survival corrected for the effect of other causes of death (Dickman et al., 2004). It represents the (hypothetical) survival that would be observed if the only possible cause of death was the disease of interest (Berkson and Gage, 1950; Schaffar et al., 2017). If we are in a situation where we have information on the cause of death, then estimation methods for net survival can be estimated in a competing risks framework. On the other hand, net survival can also be estimated when the cause of death is unknown by making use of information from the general population. In this different framework, net survival can then be estimated using the relative survival method. Since the Belgian Cancer Registry (BCR) does not collect information on the cause of death, we are in this latter situation. The net survival function is thus derived from the excess mortality hazard  $\lambda_E(t|z)$ .

Chapter 2 presents a more detailed explanation of the concepts of relative and net survival.

#### 1.2.2 Multistate models and number of years of life lost

xxx to continue.

Chapter 4 presents a more detailed explanation of the concepts of MSM and YLL.  
xxx + check all concepts covered in the thesis.



# Waiting period from diagnosis for mortgage insurance issued to cancer survivors

## 2

This chapter corresponds to an article published jointly with Catherine Legrand, Michel Denuit and Geert Silversmit, carrying the same name as the chapter and published in the *European Actuarial Journal* in 2021.

### Abstract

Massart (2018) testimonial illustrates the difficulties faced by patients having survived cancer to access mortgage insurance securing home loan. Data collected by national registries nevertheless suggest that excess mortality due to some types of cancer becomes moderate or even negligible after some waiting period. In relation to the insurance laws passed in France and more recently in Belgium creating a right to be forgotten for cancer survivors, the present study aims to determine the waiting period after which standard premium rates become applicable. Compared to the French and Belgian laws, a waiting period starting at diagnosis (as recorded in national databases) is favored over a waiting period starting at the end of the therapeutic treatment protocol. This aims to avoid disputes when a claim is filed. Since diagnosis is often recorded in the official registry database, as is the case for the Belgian Cancer Registry, its date is reliable and unquestionable in case of claim. Based on 28,994 melanoma and thyroid cancer cases recorded by the Belgian Cancer Registry, the length of the waiting period is assessed with the help of widely-accepted tools from biostatistics, including relative survival models and time-to-cure indicators. It turns out for instance that a waiting period of 4 years after diagnosis is enough for 30-year-old thyroid cancer patients. This appears to be similar to the 3-year period starting at the end of treatment protocol according to the Belgian law in such a case.

*Keywords:* Term insurance, impaired lives, cancer, home loan, right to be forgotten.

## 2.1 Introduction

Property loans are often accompanied with mortgage insurance that pays the balance of the loan if the mortgagor dies. Coverage is usually awarded in the form of term insurance with decreasing sum insured, with the amount of death benefit diminishing as the debt decreases. This is common practice in Belgium, with about 170,000 new mortgage loans per year, mainly contracted by young adults acquiring their first family house (statistics from the Belgian Central Credit Register indicate that 36% of new mortgage loans in 2017 were contracted by borrowers younger than 35 and

about 68% were granted to borrowers younger than 45).

Based on answers to a health questionnaire, insurers evaluate applicant's health status and either impose surcharges in case of impaired lives or refuse to cover the risk. Filling such health questionnaires may create frustration for patients having survived cancer occurred many years ago. Having repeatedly to answer questions related to this disease has psychological consequences and being charged higher premiums or denied coverage generates a feeling of discrimination (Massart, 2018). This is often felt as a double penalty by cancer survivors.

The restricted access to insurance cover is often regarded as a barrier to property and home ownership (in case of house loan) and to entrepreneurship (in case of professional loan). This lead Belgian authorities to create the *Bureau du suivi de la tarification assurance solde restant dû* ([www.bureaudusuivi.be](http://www.bureaudusuivi.be)) – *Opvolgingsbureau voor de tarifiering schuldsaldoverzekering* ([www.opvolgingsbureau.be](http://www.opvolgingsbureau.be)) (which could be translated literally as the “Outstanding balance insurance pricing monitoring office”) in 2014, in application of the law on insurance. This body reviews health questionnaires used by insurance companies selling mortgage insurance in Belgium and checks whether the proposed premium surcharges or cover denials are justified for impaired lives. In 2017, only 16% of the 454 cases submitted by insurance applicants to the Bureau resulted in improved policy conditions, as it can be read from the annual report published by the Bureau (2018).

Faced with a similar situation, France established in 2016 a “*droit à l'oubli*” (translated literally as “right to be forgotten” in the remainder of this text), that is, the right for an insurance applicant not to declare a previous cancer after a period of 10 years starting at the end of the therapeutic protocol. This 10-year waiting period is reduced to 5 years if the applicant suffered cancer before the age of 18. These periods of 10 and 5 years start from the date of the end of the therapeutic treatment, in absence of relapse within this period. The 10-year length of the waiting period is further shortened for several types of cancer (and other non cancer related pathologies), as detailed in the reference grid used in France (known as convention AERAS, see [www.aeras-infos.fr/cms/sites/aeras/accueil.html](http://www.aeras-infos.fr/cms/sites/aeras/accueil.html)) with reduced duration after which survivors have access to the right to be forgotten. After this period, insurance companies cannot take the pathology into account in risk assessment and cannot refuse the insurance, nor impose a premium surcharge because of the pathology.

A similar rule has been introduced in Belgium in a law dated April 4, 2019 (published in the Official Journal on April 18, 2019) for home-related and professional mortgage insurance. Despite clear similarities, the right to be forgotten established in Belgium differs from its French counterpart in an important way. Cancer survivors must still declare their pathology to insurance companies when applying for mortgage insurance in Belgium but the decision to grant coverage cannot be based on this information. Besides this right to not declare (in France) or to declare but without consequences (in Belgium) a cancer after a given waiting period, the premium surcharges may also be prohibited or limited for some cancer types. This has been implemented in Belgium in a Royal decree published in the Belgian Official Journal on June 14, 2019.

Although the establishment of such a right to be forgotten in Belgium is clearly an improvement for cancer survivors, there is most probably room for further reducing the waiting period for some cancer types. Also, there remains some ambiguity about what is considered as treatment and thus what marks the end of the therapeutic protocol. Since all cancer cases (and date of diagnosis) must be recorded in a national database in Belgium and many other EU countries, defining the start of the waiting period at the recorded date of diagnosis would certainly avoid endless discussions when a claim is filed.



There is abundant literature on cancer survival in biostatistical and medical studies. However, this topic has been the subject of few actuarial papers beyond those dealing with the assessment of extra mortality after Haberman and Renshaw (1990) and Renshaw (1988), such as Dodd et al. (2015). Let us briefly discuss some of the contributions on pricing life insurance specifically for cancer patients that appeared in the actuarial literature.

Lemaire et al. (2000) considered term life pricing in the presence of a family history of breast or ovarian cancer. These authors found that while many women with a family history of breast or ovarian cancer can be accepted at standard rates, women with two family members with cancer or one first-degree relative with cancer at an early age show substantial mortality increases (up to 100%) and can thus probably only be accepted at higher premium rates. Moreover, the authors also found that mortality increases for women with the BRCA1 or BRCA2 gene mutation (a malfunction which results in cells more likely to develop additional genetic alterations that can lead to cancer) reach 150% and can thus possibly only be accepted at a premium rate that incorporates a severe mortality surcharge.

Using 10-year and 20-year term life products, Shang (2019) calculated the single premium for breast cancer patients and found that both the average and minimum premium of the sample cancer patients to be much higher than standard premium, with the minimum premium still being close to 40% of the sum insured for the least risky patients. In order to improve the affordability of insurance products for cancer patients, Shang (2019) also suggested to set a waiting period during which no death claims will be paid. Since many deaths happen during the first years after diagnosis, a waiting period reduces the mortality risk and thus the net premium. For instance, for a 20-year term life product with sum insured 10,000 contracted by a 40-year-old cancer patient, a waiting period of 1, 2 and 3 years reduces the average premium by, respectively, 24.85%, 41.48% and 51.05% compared to no waiting period, according to the calculations by Shang (2019).

The present paper concentrates on the determination of the length of the waiting period embedded in the right to be forgotten, that is, the minimum duration before the applicant can be covered at standard premium rate. To this end, we apply several widely-accepted tools from biostatistics in order to assess excess mortality. We concentrate on melanoma (ICD-10 C43) and thyroid (ICD-10 C73.9) tumors for the sake of illustration, leaving other types of cancer for future research. These two types of cancer (called cancer sites) were selected to get a significant number of incidences occurring before the age of 40 (mortgage insurance applicants being rather young) and as they lead to a fraction of the patients who have a chance of survival close to cancer-free patients. We were thus looking for cancers with a relatively high survival rate or high cured rate. Our analysis is based on 28,994 cancer cases recorded by the Belgian Cancer Registry (BCR): 19,848 melanoma and 9,146 thyroid tumors, diagnosed between 2004 and 2016.

Based on survival data recorded by the Belgian Cancer Registry, the present paper aims

- to show that for some types of cancer (with melanoma and thyroid as examples), survivors actually have a survival comparable to that of the general population, that is, excess mortality is negligible.
- to demonstrate that patients having survived long enough to some types of cancer (still with melanoma and thyroid as examples) can access life insurance market at standard insurance rates, contrarily to the common belief within the actuarial community. The technical waiting period appears to be relatively short, and shorter compared to the 10-year period specified in the law.

In addition, we promote a waiting period starting at diagnosis rather than at the end of the therapeutic treatment protocol in order to avoid disputes in case of death. Indeed, diagnosis is recorded in databases maintained by official bodies within the European Union. The results obtained in the present study appear to be particularly encouraging as they suggest a considerable shortening of the 10-year waiting period for some types of cancer.

The remainder of this paper is structured as follows. Section 2.2 presents the data used to perform the present study. Sections 2.3 and 2.4 apply several tools from biostatistics to assess the length of the waiting period, with a focus on the estimation of the survival of cancer patients in Section 2.3 and concentrating rather on the time after which we can consider the patients still alive as “cured” in Section 2.4. Comparisons with standard premium rates based on life tables generally used on the Belgian market are provided in Section 2.5. The final section (Section 2.6) concludes the paper with a discussion.

## 2.2 Data sources

### 2.2.1 Belgian Cancer Registry (BCR)

The Belgian Cancer Registry (BCR) is a national population-based cancer registry collecting data on all new cancer diagnoses in Belgium since the incidence year 2004. For the execution of this main task, BCR relies on its own specific legislation.

In this study, we restrict our analysis to two cancer sites, as explained in the introductory section. We also limit our analyses to patients from 20 to 69 years old at time of diagnosis since the right to be forgotten mainly concerns young adults and active life. A total of 19,848 cases of melanoma and 9,146 cases of thyroid cancer diagnosed between 2004 and 2016 were followed-up until the 1<sup>st</sup> of July 2018. Follow-up thus varied from 2 years for patients diagnosed in 2016 to 14 years for those diagnosed in 2004. Patients without a national security number (INSZ/NISS) were excluded from our analyses, as we have no vital status on these patients. Patients lost to follow-up (mostly due to moving abroad) and patients still alive at the end of the follow-up period were censored. Table 2.1 summarizes the percentage of lost to follow-up before the 1<sup>st</sup> of July 2018 in each remaining subgroup together with the number of included cases. The fraction of patients lost to follow-up per subgroup varied from 1.04% for women with melanoma cancer aged 50-69 to 4.41% for male melanoma cancer patients aged 20-34. The total fraction of patients lost to follow-up cases, regardless of gender, site or age group was 1.87%. The analyses were conducted separately for women and men, representing 65.13% and 34.87% of all cases respectively, and age at diagnosis was included as a covariate, continuous or categorical depending on the approach. Age at diagnosis ranges from 20 to 69 years and 3 age groups were considered: 20-34, 35-49 and 50-69.

### 2.2.2 General population

In order to estimate excess cancer mortality, mortality in the cancer population must be compared to the expected mortality in the general population. Belgian population life tables, obtained from Statbel (the Belgian statistical office), were used to estimate expected mortality in the general population. Gender-specific mortality rates for the period 2004-2018 (by single year of age) have been smoothed in two dimensions to remove erratic variations. The surface smoothing was performed with the SAS procedure PROC LOESS using local linear polynomials weighted by population size (Cleveland et al., 1988; Cleveland and Grosse, 1991; Cleveland et al., 1992).

Gender	Cancer site	Age at diagnosis	Lost to follow-up	Number of cases included
Women	Melanoma	20-34	3.44%	1,863
		35-49	1.07%	4,386
		50-69	1.04%	5,786
	Thyroid	20-34	3.07%	1,204
		35-49	2.66%	2,596
		50-69	1.67%	3,048
Men	Melanoma	20-34	3.45%	725
		35-49	2.29%	2,360
		50-69	1.84%	4,728
	Thyroid	20-34	3.30%	273
		35-49	2.49%	724
		50-69	1.69%	1,301
Total			1.87%	28,994

**Table 2.1: Numbers of melanoma and thyroid cancer cases diagnosed in Belgium between 2004 and 2016 (BCR data) by gender, site and age group, with percentage of lost to follow-up.**

## 2.3 Survival of cancer patients

### 2.3.1 Overall survival

In this paper, we analyze survival (i.e., time to death) for cancer patients beyond diagnosis according to a number of covariates summarized into the vector  $Z$ . Specifically,  $T$  denotes the remaining lifetime at diagnosis. Given  $Z = z$ ,  $T$  has probability density function  $f(\cdot|z)$ , distribution function  $F(\cdot|z)$ , survival function  $S(\cdot|z) = 1 - F(\cdot|z)$ , and hazard rate, or force of mortality  $\lambda(\cdot|z) = f(\cdot|z)/S(\cdot|z)$ . Contrarily to insurance studies,  $T$  denotes the remaining lifetime since diagnosis and age at diagnosis is included in the covariates (attained age is thus obtained by summing age at diagnosis and survival time). This is why we refrain here from complying with the international actuarial notation for survival probabilities and force of mortality (when computing premiums in Section 2.5, we will revert back to the actuarial notation).

The non-parametric Kaplan-Meier (1958) estimator is used here to estimate the overall survival function  $S(\cdot)$ , without distinguishing according to causes of death (Belot et al., 2019). Estimated overall survival probabilities according to gender and site (melanoma and thyroid) are detailed in Table 2.2 and illustrated in Figure 2.1. The 10-year overall survival probabilities range from 0.774 (with 95% confidence interval  $[0.762 - 0.786]$ ) for men with melanoma cancer to 0.93 (with 95% CI  $[0.922 - 0.938]$ ) for women with thyroid cancer (see Table 2.2). Remember that overall survival probabilities take into account all causes of deaths, that is, both cancer and non-cancer related deaths are considered.

In addition to the two curves for women and men in Figures 2.1 and 2.2, a third curve including both sexes (denoted “Belgium”) is also drawn for the simple reason that since December 21, 2012, the European directive on equality between men and women also applies to outstanding balance insurance, so the gender no longer has an influence on the premiums. Therefore, we believe it is of interest for insurance companies to visualize these survival probabilities for both sexes combined.



**Figure 2.1: Estimated overall survival probability by gender and site using the non-parametric Kaplan-Meier (1958) estimator.**

Sex	Cancer site	$\hat{S}(t = 5)$	$CI_{95\%}$	$\hat{S}(t = 10)$	$CI_{95\%}$
Women	Melanoma	0.928	(0.923-0.933)	0.885	(0.878-0.893)
	Thyroid	0.961	(0.956-0.966)	0.930	(0.922-0.938)
Men	Melanoma	0.847	(0.838-0.855)	0.774	(0.762-0.786)
	Thyroid	0.915	(0.903-0.927)	0.839	(0.820-0.860)

**Table 2.2: Estimated overall survival probabilities by gender and site using the non-parametric Kaplan-Meier estimator at 5 and 10 years after diagnosis,  $\hat{S}(t = 5)$  and  $\hat{S}(t = 10)$ , with their confidence interval (CI) at 95% level.**

### 2.3.2 Relative survival

Information on cause of death due to cancer is often unavailable or unreliable and not all deaths can be easily classified as deaths due to the cancer of interest or due to another cause (Percy et al., 1981). Relative survival, which does not require information on the cause of death, provides a measure of the excess mortality experienced by cancer patients by comparing the mortality in the cancer population with the mortality in the general population. This led the relative survival to become the standard measure of patient survival for population-based cancer registries, as shown by its prominence around the world in studies related to cancer survival (Ries et al., 2002; Coleman et al., 1999; Berrino et al., 1999; Perme et al., 2012, 2016; Pavlič and Pohar Perme, 2019).

Relative survival models are divided into two types; (i) additive models and (ii) multiplicative models; see e.g. Pohar and Stare (2006). Despite the wide acceptance of multiplicative specifications within the actuarial community, it turns out that additive models are biologically more plausible in cancer studies and provide a better fit to the data (Dickman et al., 2004; Buckley, 1984; Hakulinen and Tenkanen, 1987; Esteve et al., 1990; Bolard et al., 2001). The additive specification is thus favored here. The hazard rate at time  $t$  since diagnosis for cancer patients with covariate vector  $\mathbf{Z}$ , is decomposed into two additive components: the population hazard based on available patient's characteristics  $\mathbf{Z} = \mathbf{z}$ , denoted as  $\lambda_P(\cdot|\mathbf{z})$ , and the excess hazard specific for the disease of interest, denoted as  $\lambda_E(\cdot|\mathbf{z})$ . Formally,

$$\lambda(t|\mathbf{z}) = \lambda_P(t|\mathbf{z}) + \lambda_E(t|\mathbf{z}). \quad (2.1)$$

From this expression, relative survival model (2.1) can be written as

$$S(t|\mathbf{z}) = S_P(t|\mathbf{z})r(t|\mathbf{z})$$

where the relative survival function  $r(\cdot|\mathbf{z})$  is defined as

$$r(t|\mathbf{z}) = \frac{S(t|\mathbf{z})}{S_P(t|\mathbf{z})}. \quad (2.2)$$

In words, the relative survival function  $r(\cdot|\mathbf{z})$  corresponds to the ratio of the survival function of the studied group  $S(\cdot|\mathbf{z})$  to the survival function of a comparable group (i.e., with the same characteristics) from the general population  $S_P(\cdot|\mathbf{z})$  (Dickman et al., 2004).

In (2.1),  $\lambda_P(\cdot|\mathbf{z})$  is usually estimated on the basis of external data such as population life tables, which are usually stratified according to the main factors affecting patient survival such as age, gender, and calendar year. As population life tables take into account all deaths, those due to the cancer of interest are thus also included. However, it is assumed that this does not influence the estimated ratio as mortality for a given cancer represents only a small fraction of the overall mortality. Correcting for this over-representation of the cancer being studied has, in practice, an insignificant effect on estimates of expected survival (Esteve et al., 1994). Oksanen (1998) shows that this holds even for common cancers such as prostate cancer.

Net survival is a measure of patient survival corrected for the effect of other causes of death (Dickman et al., 2004). It represents the (hypothetical) survival that would be observed if the only possible cause of death was the disease of interest (Berkson and Gage, 1950; Schaffar et al., 2017). If we are in a situation where we only have data from (a sample of) our sub-population of interest (e.g., cancer patients) but we have information on the cause of death, then estimation methods for net survival, sometimes referred then as marginal survival (see Geskus, 2015), can be estimated in a competing risks framework. However, unbiased estimate of this net or marginal (cause-specific) survival can only be obtained if one can assume independence of the



**Figure 2.2: Net survival by gender and site using the non-parametric Perme et al. (2012) estimator.**

censoring due to death from other causes (Latouche et al., 2013; Schaffar et al., 2017). On the other hand, net survival can also be estimated when the cause of death is unknown by making use of information from the general population. In this different framework, net survival can then be estimated using the relative survival method. Since the BCR does not collect information on the cause of death, we are in this latter situation. The net survival function is thus derived from the excess mortality hazard  $\lambda_E(\cdot|z)$ . Formally, it is defined as

$$S_n(t|z) = \exp\left(-\int_0^t \lambda_E(u|z)du\right). \quad (2.3)$$

Depending on  $\lambda_E(\cdot|z)$ ,  $S_n(\cdot|z)$  may be a proper survival function but this is not necessarily the case.

There are several approaches to estimate net survival of a cohort of patients in a relative survival framework. Danieli et al. (2012) showed that only two of them provide unbiased estimates of net survival: (i) the non-parametric Perme et al. (2012) estimator and (ii) the excess risk based on an adjusted modeling on the demographic variables of the life tables. We used the non-parametric Perme et al. (2012) estimator to estimate net survival as recommended by Danieli et al. (2012) for population based studies.

Net survival probabilities by gender and site are illustrated in Figure 2.2, while net survival by age group and site is displayed in Figure 2.3. Numerical values are listed in Table 2.3. The net survival functions reach a plateau for both sites and genders, which is indicative for ‘cure’ of cancer. The estimated survival curves staying practically constant after 5 years since diagnosis indicates that the excess hazard of dying compared to the general population becomes negligible after only a few years after diagnosis. This suggests that a waiting period of moderate length would be enough to apply standard life insurance rates for the cancers under consideration.

Net survival by site and age group (both sexes combined, Figure 2.3) follows a quite expected path for melanoma cancer patients (left panel): younger patients have a better net survival compared to older patients. In particular, the 5-year and 10-year net survival probabilities for patients aged 20-34 were respectively 0.959



**Figure 2.3: Net survival by age group and site using the non-parametric Perme et al. (2012) estimator.**

Gender	Cancer site	Age at diag.	$\hat{S}_n(t = 5)$	$CI_{95\%}$	$\hat{S}_n(t = 10)$	$CI_{95\%}$
Women	Melanoma	20-34	0.974	(0.966-0.982)	0.963	(0.951-0.974)
		35-49	0.956	(0.949-0.963)	0.936	(0.925-0.946)
		50-69	0.927	(0.918-0.936)	0.907	(0.893-0.922)
	Thyroid	20-34	0.998	(0.995-1.000)	0.997	(0.990-1.000)
		35-49	0.992	(0.987-0.997)	0.990	(0.981-0.998)
		50-69	0.958	(0.948-0.968)	0.941	(0.924-0.959)
Men	Melanoma	20-34	0.921	(0.901-0.942)	0.894	(0.868-0.921)
		35-49	0.899	(0.886-0.913)	0.871	(0.853-0.890)
		50-69	0.871	(0.859-0.884)	0.849	(0.828-0.870)
	Thyroid	20-34	0.997	(0.987-1.010)	0.989	(0.968-1.010)
		35-49	0.972	(0.957-0.987)	0.932	(0.901-0.964)
		50-69	0.928	(0.908-0.949)	0.895	(0.858-0.933)

**Table 2.3: Net survival probabilities by gender, site and age group using the non-parametric Perme et al. (2012) estimator at 5 and 10 years after diagnosis,  $\hat{S}_n(t = 5)$  and  $\hat{S}_n(t = 10)$ , with their confidence intervals at 95% level.**

(with 95% confidence interval  $[0.951 - 0.968]$ ) and 0.943 (with 95% confidence interval  $[0.932 - 0.955]$ ), meaning that for this age group (which is typically the age at which one starts a loan), patients' survival is close to that of the general population. Net survival by age group for thyroid cancer patients (right panel of Figure 2.3) yields even more promising results in the context of mortgage loans as the age group which is most likely to subscribe to such financial products (20-34 years) has a 5-year and 10-year net survival probabilities of respectively 0.998 (with 95% confidence interval  $[0.995 - 1]$ ) and 0.996 (with 95% confidence interval  $[0.989 - 1]$ ).

Net survival by gender, site and age group (Table 2.3) indicates a highly favorable outcome for women and men diagnosed with thyroid cancer and aged 20-34. For both subgroups, the 95% confidence interval for net survival at 10 years after diagnosis indicate that a net survival equal to that of the general population cannot be rejected at the 5% significance level.

Regarding the gap between women and men for the net survival curves (in particular for melanoma cancer patients), we do not have a clear explanation for these differences in survival probability. However, although stage distribution is very similar across gender, The Belgian Cancer Registry (2012), in agreement with Balch et al. (2001), showed that compared to men, females have more often melanoma on the arms or legs which has a better prognosis. This partly explains the difference in survival for melanoma cancer patients.

Using data provided by the French network of cancer registries (FRANCIM), Boussari et al. (2018) found very similar results for thyroid cancer patients diagnosed between 1995 and 2010. They obtained a 10-year net survival of 0.99 (95% CI  $[0.99-1.00]$ ) for women aged 15-45, and 0.98 (95% CI  $[0.96-0.99]$ ) for men of the same age group. Their results are also very close to ours when comparing patients aged 45-55: a 10-year net survival of 0.99 (95% CI  $[0.98-1.00]$ ) for women, and 0.92 (95% CI  $[0.88-0.95]$ ) for men. In another study including melanoma cancer patients diagnosed between 1989 and 2004 by the French registries, Jooste et al. (2013) obtained a 10-year net survival for men and women aged 15-45 of 0.81 (95% CI  $[0.78-0.84]$ ) and 0.91 (95% CI  $[0.89-0.93]$ ), respectively.

### 2.3.3 Proportional excess hazards

Previous sections suggested that some patients actually have a survival comparable to that of the general population. In the following sections, we confirm these findings from the point of view of the excess hazard.

Esteve et al. (1990) proposed a maximum likelihood method for estimating net survival via the modeling of the excess hazard. The excess hazard  $\lambda_E(t|z)$  to be estimated is represented as

$$\log(\lambda_E(t|z)) = (\beta^\top z) + \log \left( \sum_{k=1}^m \tau_k I_k(t) \right) \quad (2.4)$$

where  $\beta$  is the log hazard ratio corresponding to the covariates,  $I_k(t)$  the indicator function for the  $k^{th}$  interval (after splitting the follow-up time into short time intervals) and  $\tau_k$  the net baseline hazard rate in that interval for patients with  $z = \mathbf{0}$ .

A maximum likelihood approach to estimate parameters of model (2.4) is available in the `relsurv` package (Perme and Pavlič, 2018; Pohar and Stare, 2006) in R (R Core Team, 2017). Fitting two models to our data (one for each cancer site) with age group (50-69 years old taken as the reference category) and follow-up (with intervals of 5 years) as covariates, we obtain the results presented in Tables 2.4 and 2.5. The coefficient estimates  $\hat{\beta}$  and their standard errors are displayed in the second column,  $p$ -values are reported in the last column. Note that the significance is largely impacted by the large sample size.



Covariates	$\hat{\beta}$ (S.E.)	$p$ -value
Gender Women	-0.759 (0.060)	<0.001
Agegr 20-34	-0.741 (0.105)	<0.001
Agegr 35-49	-0.349 (0.064)	<0.001
fu [0,5)	-3.507 (0.045)	<0.001
fu [5,10)	-4.590 (0.105)	<0.001
fu [10,14]	-5.334 (0.420)	<0.001

**Table 2.4: Results of model (2.4) fitted to melanoma cancer data.**

Covariates	$\hat{\beta}$ (S.E.)	$p$ -value
Gender Women	-0.754 (0.169)	<0.001
Agegr 20-34	-2.864 (0.703)	<0.001
Agegr 35-49	-1.252 (0.213)	<0.001
fu [0,5)	-4.086 (0.129)	<0.001
fu [5,10)	-5.089 (0.314)	<0.001
fu [10,14]	-5.276 (0.867)	<0.001

**Table 2.5: Results of model (2.4) fitted to thyroid cancer data.**

For both cancer sites, being a woman and being younger at the time of diagnosis are good prognostic factors.

#### 2.3.4 Flexible parametric model

Model (2.4) is based on the assumption of proportional excess hazards, which constrains the hazard ratio to be constant over the follow-up time (Giorgi et al., 2005). Nonetheless, in cancer survival, the effects of prognostic factors often vary with time since diagnosis (Dickman et al., 2004; Quantin et al., 1999) and it is well known that the linearity assumption of covariates may be too strong and not always verified in practice (Mounier, 2015). As an example with the variable age, the effect on hazards of an increase of one year is often different for patients aged 18 or 60. This is why Remontet et al. (2019) and Fauvernier et al. (2019a,b) extended the model proposed by Esteve et al. (1990) to account for non-linear and non-proportional effects of covariates. This model also allows (i) a flexible modeling of the baseline hazard and (ii) a flexible interaction between several covariates adopting a multidimensional penalized splines approach. This leads to the specification

$$\log(\lambda_E(t|z)) = \sum_{j=1}^J g_j(t, z) \quad (2.5)$$

where  $g_j(\cdot, \cdot)$  are uni- or multidimensional penalized spline functions and each function  $g_j(\cdot, \cdot)$  can be the marginal basis of time, the marginal basis of a covariate or the tensor product of the marginal bases of any number of elements of  $(t, z)$  (Fauvernier et al., 2019b). This model has the advantage that the splines bring the flexibility needed for modeling the hazard and inclusion of penalty terms allows to control this flexibility for smooth estimation (as suggested by Eilers and Marx, 1996).

Several flexible models to estimate excess hazard were considered in this paper:

- baseline hazard only (BH) model:  
Eq. (2.4) without the  $(\beta^\top z)$  term, as it considers an excess hazard in each interval and no other covariates



**Figure 2.4: Excess hazard by age and cancer site estimated with a non-linear and non-proportional hazard model.**

- linear and proportional hazard (LPH) model:  
 $\log(\lambda_E(t|z)) = f(t) + age$
- linear and non-proportional hazard (LNPH) model:  
 $\log(\lambda_E(t|z)) = f(t) + age + g(t) \cdot age$
- non-linear and proportional hazard (NLPH) model:  
 $\log(\lambda_E(t|z)) = f(t) + g(age)$
- non-linear and non-proportional hazard (NLNPH) model:  
 $\log(\lambda_E(t|z)) = f(t) + g(age) + g(t) \cdot age$

with  $f(t)$  the flexible parametric function for the baseline/reference hazard as a function of time and  $g(z)$  the (non-)linear function of the covariates.

We fitted all these models and compared them based on a likelihood ratio test. The remaining of this section is based on the NLNPH model, deemed the best one according to likelihood ratio test. Excess hazard, assuming non-linear and non-proportional hazard for age at diagnosis, are estimated using the `flexrsurv` package (Clerc-Urmès et al., 2020) in R.

As previously suggested, for both cancer sites, excess mortality hazard increases with age at diagnosis and decreases with time since diagnosis (Figure 2.4). For thyroid cancer patients, excess hazard at the time of diagnosis is approximately 0.15 for patients aged 65 and is close to 0 for patients aged 40 and below. From 4 years to 10 years after diagnosis, excess hazard remains constant and is close to 0 for all ages. Beyond 10 years after diagnosis, excess hazard slightly bends up but remains small. This observation can also be an artefact of the spline function, which is “unbounded” at the end (a higher order degree for the spline function can produce this upwards bend). For melanoma cancer patients, excess hazard at the time of diagnosis is approximately 0.035 for patients aged 65 and is close to 0 for the youngest patients. For all ages, excess hazard peaks at 1 year after diagnosis before decreasing until it becomes negligible, around 8 years after diagnosis.

### 2.3.5 Cure models

Cure models are a specific class of survival models which assume that a fraction of the subjects will never develop the event of interest, here death due to cancer. Such models have been used in different fields such as economics (e.g., time until an unemployed person finds a new job), engineering (e.g., time until a machine or device fails), finance (e.g., time until a bank goes bankrupt), marketing (e.g., time until a client buys a new product), for instance. In our context, cure models can be used to determine cancer patients who are considered as “long-term survivors” and those who are not (Maller and Zhou, 1996; Othus et al., 2012). The long-term survivors are still often referred to as “cured subjects” in the literature as a consequence of the name of this class of models. Cure models are particularly suitable for some cancer sites because if the treatment is successful, the patient will never suffer a relapse of the disease. It is also particularly suitable for less aggressive cancers such as, among others, the two considered in this paper because a considerable percentage of subjects exhibits long-term survival.

There are two main types of cure models in the literature: (i) mixture cure models (the most common type, based on the seminal work by Boag (1949) and Berkson and Gage (1952)) and (ii) non-mixture cure models (Andrei et al., 1996; Tsodikov et al., 2003; Chen et al., 1999); see Amico and Van Keilegom (2018) for a recent overview. In the present paper, we consider only the family of mixture cure models. In this approach, the patient population is considered as a mix of two types of patients, that is, long-term survivors who will never die of their cancer and the uncured patients who, if not censored, will die of their cancer (Lambert et al., 2006). In the global survival setting, the mixture cure model is specified as follows:

$$S(t) = \pi + (1 - \pi)S_u(t),$$

where  $\pi$  is the proportion of patients that are long-term survivors and  $S_u(\cdot)$  the survival function of the uncured population. Both  $\pi$  and  $S_u(\cdot)$  can then be modeled to depend on covariates. Cure models can be a useful alternative to standard survival models for cancers with a strong medical evidence and a confirmation in the data for the presence of long-term survivors (Legrand and Bertrand, 2019). In the relative survival setting, cure models also allow to determine the proportion of statistically cured cases and survival time of the fatal cases (Silversmit et al., 2017).

The estimated proportion of cured cases and mean survival time of fatal cases for the two cancer sites considered (melanoma and thyroid), using a mixture cure model are given in Table 2.6. Note that NAs for young age groups are due to an insufficient number of cases.

The estimated cured proportion ranges from 84.99% ( $CI_{95\%}$  [84.65, 85.33]) for men with melanoma cancer aged 50-69 to 99.86% ( $CI_{95\%}$  [99.82, 99.90]) for women with thyroid cancer aged 20-34. The estimated mean survival time of fatal cases ranges from 0.47 years ( $CI_{95\%}$  [-0.18, 1.12]) for women with thyroid cancer aged 20-34 to 10.57 years ( $CI_{95\%}$  [1.32, 19.81]) for men with thyroid cancer aged 35-49. Moreover, higher age at diagnosis is correlated with lower cured proportions, except for men with thyroid cancer from the oldest age group.

Using data on 818,902 Italian cancer patients diagnosed between 1985 and 2005, Dal Maso et al. (2014) also found encouraging results for thyroid cancer patients, with an estimated cured fraction of 99% and 95% for women and men aged 15-45, respectively. Regarding melanoma cancer patients, the results on Italian data are somewhat lower than the estimations presented in this work, with a cured proportion of 85% and 77% for women and men aged 15-45, respectively. The fact that we are using a more recent incidence period may partly explain these improved results.

Gender	Cancer site	Age at diag.	Est. cured fraction	$CI_{95\%}$	Est. mean $T$	$CI_{95\%}$
Women	Melanoma	20-34	96.53	(96.33-96.72)	3.81	(3.46-4.22)
		35-49	93.28	(92.94-93.63)	4.49	(4.04-5.03)
		50-69	90.01	(89.49-90.53)	3.93	(3.50-4.46)
	Thyroid	20-34	99.86	(99.82-99.90)	0.47	(-0.18-1.12)
		35-49	99.03	(98.82-99.23)	4.87	(3.03-6.71)
		50-69	95.73	(95.39-96.07)	1.40	(1.02-1.78)
Men	Melanoma	20-34	88.77	(88.21-89.32)	3.54	(3.12-4.07)
		35-49	86.28	(85.77-86.78)	3.83	(3.52-4.20)
		50-69	84.99	(84.65-85.33)	2.94	(2.79-3.10)
	Thyroid	20-34	NA	NA	NA	NA
		35-49	91.96	(86.99-96.93)	10.57	(1.32-19.81)
		50-69	92.87	(92.34-93.41)	0.96	(0.67-1.25)

**Table 2.6: Estimated cured fractions (in %) and mean survival time (in year) for the fatal cases by gender, site and age group, with their 95% confidence intervals. *Note: Est. mean  $T$  = Estimated mean survival time of fatal cases. NAs for young age groups are due to an insufficient number of cases.***

## 2.4 Time-to-cure

The time-to-cure (TTC) is generally referred as the time after which patients can be considered as long-term survivors. Different approaches to define TTC have emerged in the literature. We highlight three of them. Firstly, Chauvenet et al. (2009) defined  $TTC_C$  as the time at which “almost” (that is,  $1 - \epsilon$ , with  $\epsilon$  small enough usually ranging from 0.1 to 0.01) all uncured patients would have died. From that time onwards, the number of deaths attributable to the cancer of interest becomes negligible.

Secondly, Dal Maso et al. (2014) define  $TTC_D$  as the shortest time after diagnosis at which the 5-year conditional net survival (defined as the ratio between net survival at time  $t + 5$  years and net survival at time  $t$ ) is close to 1.

Thirdly, Boussari et al. (2018) defined  $TTC_B$  as the shortest time from which the conditional probability of being cured at a given time  $t$  after diagnosis knowing that the patient was alive up to time  $t$  is close to 1, that is  $TTC_B$  is the smallest value of  $t$  such that, for some given (small) value of  $\epsilon$

$$\frac{\pi}{S(t)} = \frac{\pi}{\pi + (1 - \pi)S_u(t)} \geq 1 - \epsilon, \quad (2.6)$$

where  $\pi$  is the proportion of cured patients and is estimated from the relative survival with the hypothesis of cure.

The main advantage of TTC is that, although the results may depend on the chosen definition, it is a simple and straightforward indicator to set the time after which a patient who had cancer should not be penalized anymore when applying for mortgage insurance.

In this paper, we focus on  $TTC_B$  for several reasons. First,  $TTC_B$  depends on both the cure proportion and the survival of the uncured, so it is less influenced by high early excess mortality. Second,  $TTC_B$  has the advantage of being an increasing function of time, therefore, it is not sensitive to a temporary plateau effect (that is, net survival curve flattening before decreasing again). Third, estimating  $TTC_B$  requires a 5-year shorter follow-up than  $TTC_D$ , which is a clear advantage of  $TTC_B$  over  $TTC_D$  (Boussari et al., 2018).

Time-to-cure  $TTC_B$  was estimated using the `rstpm2` package (Clements and Liu,

Gender	Cancer site	Age at diag.	$\widehat{TTC}_B$ $\epsilon = 0.05$	$CI_{95\%}$ $\epsilon = 0.05$	$\widehat{TTC}_B$ $\epsilon = 0.01$	$CI_{95\%}$ $\epsilon = 0.01$	Est. cure prop. (%)
Women	Melanoma	20-34	0.01	(0.00-1.18)	4.32	(3.30-5.35)	96.75
		35-49	0.81	(0.21-1.42)	5.65	(4.82-6.49)	94.16
		50-69	2.56	(2.04-3.08)	6.39	(5.66-7.12)	90.87
	Thyroid	20-34	0.01	(0.00-23.42)	0.01	(0.00-23.42)	99.75
		35-49	0.01	(0.00-14.45)	0.01	(0.00-14.45)	99.35
		50-69	0.01	(0.00-5.10)	4.82	(3.78-5.86)	94.93
Men	Melanoma	20-34	2.38	(1.92-2.85)	5.64	(4.90-6.39)	89.69
		35-49	2.78	(2.33-3.23)	5.86	(5.13-6.59)	88.12
		50-69	3.42	(3.00-3.85)	6.21	(5.51-6.90)	84.77
	Thyroid	20-34	0.01	(0.00-28.97)	0.11	(0.00-0.78)	98.88
		35-49	0.01	(0.00-16.03)	4.66	(3.19-6.13)	96.46
		50-69	0.91	(0.51-1.31)	7.14	(6.40-7.87)	90.84

**Table 2.7: Estimated value of time-to-cure ( $\widehat{TTC}_B$  in years, with  $\epsilon = 0.05$  and  $\epsilon = 0.01$ ) together with 95% confidence intervals and cure proportion (in %) by cancer site, sex and age group.**

2019) in R. Table 2.7 presents estimated  $\widehat{TTC}_B$  (with  $\epsilon = 0.05$  and  $\epsilon = 0.01$ ) in years and the cure proportion in percentage for each subgroup.

In the following we interpret only  $\widehat{TTC}_B$  with  $\epsilon = 0.05$ . Results of  $\widehat{TTC}_B$  with  $\epsilon = 0.01$  are still presented for the sake of comparison and completeness. From Table 2.7 we see that while  $\widehat{TTC}_B$  increases with age, ranging from a few days (0.01 year) for the youngest age groups to almost 3.5 years for the oldest age group of male melanoma cancer patients, cure proportion decreases with age, ranging from 99.75% for women aged 20-34 with thyroid cancer to 84.77% for men aged 50-69 with melanoma cancer. With this approach, the subgroups that stand out the most are the ones with a small  $\widehat{TTC}_B$  (and confidence intervals as narrow as possible to decrease uncertainty as much as possible) and a large proportion of long-term survivors. Among all subgroups considered, this is the case especially for women with melanoma cancer aged 20-34 ( $\widehat{TTC}_B = 0.01$  with 95% confidence interval [0.00–1.18] and estimated cure proportion = 96.75%).

Figure 2.5 illustrates the cure proportion obtained from a cure model and the  $\widehat{TTC}_B$  ( $\epsilon = 0.05$ ) by age group and gender for both cancer sites. We can easily classify points into two clusters; the ones in the upper left corner with the best possible outcomes in terms of  $\widehat{TTC}_B$  and cure proportion and the others. Among melanoma cancer patients, women aged under 50 belong to this group with favorable outcomes, whereas for thyroid cancer patients, women of all ages and men aged under 50 belong to this group.

Similar results for thyroid cancer patients have been obtained by Boussari et al. (2018) on FRANCIM data. They also found that women aged 15-65 and men aged 15-45 have highly favorable outcomes, that is, a cure proportion close to 100% and a  $\widehat{TTC}_B$  close to 0.

## 2.5 Application to mortgage insurance

All results obtained so far suggest that, for melanoma and thyroid cancer patients, excess mortality becomes negligible after some waiting period. In this section, we determine the length of such a waiting period as the time needed to get back to standard premium rates. Henceforth, standard rates correspond to premiums computed according to life tables commonly used on the Belgian market:

- regulatory life table XK applying to insurance products comprising benefits in



**Figure 2.5: Cure proportion versus time-to-cure (in years, with  $\epsilon = 0.05$ ) by gender and age group at diagnosis in patients diagnosed with melanoma and thyroid cancer.**

case of death (formally, XK defines minimum premium amount for policies with a positive sum at risk). This life table is conservative and generates a relatively high safety loading.

- experience market life table published by the National Bank of Belgium (NBB). These life tables reflect the mortality observed on the market, within portfolios of companies controlled by NBB. There is no safety loading and insurers are only allowed to apply premium rates resulting from NBB tables for relatively short periods of time (rates are subject to revision in case the observed mortality on the market changes over time).

We also include premium rates calculated from general population life table published by Statbel, for the sake of comparison. Term life premiums are smaller for NBB life tables and larger for XK life table, Statbel life tables falling in between because of the higher socio-economic profile of the insured population. Premium rates for cancer patients are computed from the excess mortality hazards estimated with the help of the flexible parametric model discussed in Section 2.3.4, according to the time elapsed since diagnosis.

Consider a mortgage insurance applicant aged  $x$  borrowing an amount of 100,000 at interest rate 2% for a duration  $\delta$ . At time  $t$ , the amount of the loan that has not been amortized is denoted as  $c(t)$ . This loan is secured by mortgage insurance, repaying the lender the amount  $c(t)$  in case the policyholder dies at time  $t$ . The expected present value (EPV) of benefits paid in case of death is thus equal to

$$\text{EPV} = \int_0^{\delta} c(t)v(0, t)_t p_x \mu_{x+t} dt \quad (2.7)$$

where  $c(t)$  is the amount of benefit in case of death at time  $t$  (which is in our case, a decreasing sum insured corresponding to the amount of the loan not yet amortized),  $v(0, t)$  is the present value at time 0 of a unit payment made at time  $t$  (the discount factor),  $_t p_x$  is the  $t$ -year survival probability for a policyholder aged  $x$  and  $\mu_{x+t}$  is the force of mortality at attained age  $x + t$ .



**Figure 2.6: Expected present value (EPV) of a life insurance contracted by a 30 and 50-year-old cancer patient for a period of 20 years with interest of 1 percent and benefit of 100,000. Horizontal lines correspond to EPV calculated with XK, NBB and Statbel life tables.**

In case the applicant suffered from cancer, we have to relate  ${}_t p_x$  and  $\mu_{x+t}$  entering the formula to the survival function and hazard rate estimated in the preceding sections. To this end, we assume that the applicant aged  $x$  has been diagnosed with cancer at age  $x - w$ . We then have

$$\begin{aligned} {}_t p_x &= \frac{S(w + t | \text{age at diagnosis} = x - w)}{S(w | \text{age at diagnosis} = x - w)} \\ \mu_{x+t} &= \lambda(w + t | \text{age at diagnosis} = x - w). \end{aligned}$$

with  $S(\cdot)$  and  $\lambda(\cdot)$  being estimated from the NLNPH model (eq (2.5)) discussed in Section 2.3.4.

This allows us to compute  $EPV(w)$  according to (2.7) for each candidate waiting period  $w$  after diagnosis and to select the smallest  $w$  such that  $EPV(w)$  becomes close to the value computed from the XK life table (assuming that cancer patients are priced with the regulatory life table XK and that the market can absorb the extra mortality burden corresponding to the difference between XK and NBB life tables).

Consider a home loan of duration 20 years. A cancer patient aged 30 and another aged 50 apply for mortgage insurance, with technical interest rate of 1% and a term of 20 years. These characteristics have been chosen as they represent a rather standard setting and other scenarios revealed similar patterns when considering different ages between 18 and 50. EPV based on XK, NBB and Statbel life tables have been computed to compare it with  $EPV(w)$  for a cancer patient diagnosed with melanoma or thyroid cancer at ages  $30 - w$  and  $50 - w$ . Results are illustrated in Figure 2.6. Notice that EPV are presented for both genders combined, as XK life table applies to both sexes and insurance companies operating in the European Union are not allowed to account for gender in pricing.

Results show that EPV of a 30-year-old patient approaches EPV based on XK life table about 9 years after diagnosis for melanoma cancer patients and after slightly less than 1 year for thyroid cancer patients. For a 30-year-old patient with melanoma cancer, the EPV never reaches within our time period the EPV based on NBB and

Statbel life tables. For a 30-year-old patient with thyroid cancer, the EPV goes below the lowest EPV (based on NBB life table) as early as about 4 years after diagnosis. For a 50-year-old patient with melanoma cancer, EPV reaches the same level than the one based on XK life table 3 years after diagnosis, and reaches the lowest level (based on NBB life table) less than 8 years after diagnosis. Finally, for a 50-year-old patient with thyroid cancer, EPV reaches XK level 1 year after diagnosis, then stays below the lowest EPV 3 years after diagnosis. This is in line with the reduced waiting periods published in the Royal decree on June 14, 2019, where a duration of 3 years applies in this case. There is however a fundamental difference in the approach because the waiting period starts at diagnosis in the present study whereas it starts at the end of the treatment protocol according to the law.

The improvement for melanoma cancer patients is not as substantial as for thyroid cancer patients since the time period is close to 9 and 3 years for 30 and 50-year-old patients, respectively. Nonetheless, it remains advantageous for patients since the 9 and 3-year period starts from the date of diagnosis and not at the end of the therapeutic treatment. Note that, at first glance there seems to be an advantage for older patients with shorter waiting time, while all the other results seem rather to indicate an advantage for younger patients with shorter time to cure and higher cure rate. This actually comes from the fact that in absolute terms, younger patients have shorter time to cure and thus lower EPV than older patients. As shown in Figure 2.6, EPV for a 30-year-old patient is lower than for a 50-year-old patient regardless of the time since diagnosis and for both cancer sites. However, financial burdens (i.e., XK, NBB and Statbel levels) are much lower for younger people than for older people (since young people from the general population have a lower probability of dying than older people).

## 2.6 Discussion

Results derived in this paper are in line with the reduced waiting period specified in the Belgian legislation. Furthermore, results are also in line with the reference grid used in France (convention AERAS) as the time after which patients have access to the right to be forgotten according to this convention is relatively short (maximum 6 years after the end of the therapeutic protocol for the two cancers considered in this paper).

All analyzes in this paper are based on the time since diagnosis although the right to be forgotten implemented in Belgium and France is applicable after a certain time after the end of the therapeutic protocol. For the sake of clarity and easiness, an approach based on the time since diagnosis would undeniably be more favorable to patients. A right to be forgotten based on the date of diagnosis would indeed allow patients to know when exactly they can expect to benefit from this right. On the contrary, with the current approach based on treatment end date (which is unknown until the success of the treatment and can even change later in case of relapse), benefiting from this right is subject to a high level of uncertainty as durations of treatments are heterogeneous and unpredictable even within same cancer types and stages. Therefore, patients cannot currently easily estimate when (and if) they will be able to benefit from this right.

One of the main results of this paper is thus to promote the use of the date of diagnosis instead of the end of the therapeutic treatment for defining the waiting period. However, one could argue that the length of medical treatments may have significantly reduced over the period 2004-2018 and if this is the case, both approaches could be now closer than expected. While it could have been interesting to formally compare both approaches, individual data on the type and length of treatment for each case is not reported in the Belgian Cancer Registry and such information is



not readily available. Furthermore, the definition of the end of the treatment is in itself debatable and the duration of the treatment can be quite different depending on several factors, which is actually an argument in favor of considering the date of diagnosis. Moreover, durations of treatment are heterogeneous even within the same cancer type, usually unpredictable, and optimal durations are often still open to debates (Schvartsman et al., 2019). In any case, a reduction in treatment length due to the progress made in medical treatment of cancer would obviously lead to closer agreement between the two approaches. Since the date of diagnosis, as recorded in national registries, offers the advantage to not be subject to any discussion and to allow the patient to know from the start when the waiting period will end, we think that all parties (actuaries included) will benefit from using the date of diagnosis instead of the end of treatment for more convenience and less uncertainty.

Contrarily to other studies (like Yue et al., 2018), calendar time has not been included in the analysis conducted in the present paper, because of the limited amount of cases available (recall that we concentrate on younger ages because of the product under consideration, targeting young adults). That being said, we performed two subanalyses: one for the cohort 2004-2011 and one for the cohort 2012-2018 and compared the results. There are no real changes, so we conclude that there is no cohort effect. Moreover, simplicity of the system is crucial and given that medical treatments keep improving, the resulting bias of ignoring a potential cohort effect favors insurance providers. Notice that the approaches are dependent on factors such as changing diagnostic criteria and improved diagnostic methods. As these factors may vary over time irrespective of any improvement of the treatment and are different between populations, one can not compare excess risks across different time periods and populations (Lenner, 1990). To illustrate this, suppose that a medical advance allows a cancer to be diagnosed at a less severe stage (cases that are not as fatal as the ones detected with the previous methods) and perhaps also earlier with the consequence that more cases are detected (cases that would not have been detected with the previous detection methods are now detectable). These improvements will yield an increased survival rate, regardless of whether the treatment improved or not. This weakness of the survival rate has been pointed out in the literature extensively (Enstrom and Austin, 1977; Bailer III and Smith, 1986, among others). However, although not necessarily the case, earlier diagnoses will in most cases be associated with better efficacy of the treatment.

The melanoma and thyroid cancers may include a variety of types and could be diagnosed at different stages of severity. Moreover, significant gaps are observed between women and men. It is undeniable that including the information on stages of severity and gender in the NLNPH model would refine the analysis. However, these have been ignored on purpose, considering the specific application of the results for insurance practice. Since December 21, 2012, the European directive on equality between men and women also applies to outstanding balance insurance. The judgment of the European Court of Justice indeed stipulates that no discrimination can be made between men and women when establishing such insurance contract, so the gender no longer has an influence on the premiums nor on the coverage conditions of outstanding balance insurances. Therefore, the gender has been omitted not only for the sake of simplicity, but also because it is illegal for insurance companies to use that information. Concerning the stage of severity, it has also been omitted to ensure the simplicity and thus the legal safety of the system. Moreover, not including the stage may bias the analysis but not necessarily in favor of patients diagnosed at an advanced stage (perhaps a shorter waiting period would be indicated for them). In any case, if the method amounts to determining the time to wait before mortality returns to normal, ignoring the stage is actually a conservative approach for the insurer.

To ensure that the coverage cost of cancer patients remains acceptable for the

insurance industry, further constraints may be imposed in terms of sum insured, for instance. Also, it could be reasonable to impose that insurers charge a single premium for mortgage insurance to mitigate mortality risk. Last but not least, compensation could be performed at market level to avoid that some insurers face higher costs. Indeed, even if the coverage of mortality risk becomes affordable after a relatively short waiting period, premium rates reflecting the actual mortality of cancer survivors remain sometimes above the market premiums resulting from NBB life tables. This means that when it is the case, this extra cost must be fairly distributed among stakeholders: cancer patients, insurance industry, banking sector (as they sell the loans) and society as a whole.

Notice that cancer stage at diagnosis has not been taken into account in the present study. For patients with more advanced cancer stage, the waiting period will be more conservative because mortality will peak just after diagnosis and before reverting back to the general population level. Note also that *in situ* cancer cases (considered as pre-cancer) have not been included in the present study as they are not classified as cancer *per se* (Chang et al., 2007) like any other “regular” cancer cases which were not diagnosed as *in situ* before. Moreover, cancer is not one disease, but a family of many diverse diseases with different outcomes. Results in the present paper focus on melanoma and thyroid cancer patients, and cannot be applied to other cancer types. A natural extension of this work would be to repeat the analyses for all major cancer types. This would certainly be useful for implementing appropriate market rules but goes beyond the scope of this study which primarily aims to advocate a waiting period starting at diagnosis.

Cancer patient survival has improved over the last few decades, with an increasing proportion of patients being cured for many types of cancer (Andersson et al., 2011; Lambert et al., 2006). Providing coverage in case cancer is diagnosed or to long-term cancer survivors is therefore of prime importance, for the society but also for the insurance industry since proper coverage of such risks may well produce attractive returns.

# Semi-Markov modeling for cancer insurance

3

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## Abstract

Advancements in medicine and biostatistics have already resulted in a better access to insurance for people diagnosed with cancer. This materializes into the “right to be forgotten” adopted in several EU member states, granting access to insurance after a waiting period of at most 10 years starting at the end of the successful therapeutic protocol. This paper concentrates on insurance covers on a market where such a right has been implemented. Stand-alone products are considered, as well as guarantees included as a rider in an existing package. The cost of offering standard premium rates to all applicants in mortgage insurance related to property loans is also evaluated. The 3-state (healthy–ill–dead) Semi-Markov hierarchical model developed in Denuit et al. (2019) for long-term care insurance is adopted here for actuarial calculations. Semi-Markov transition intensities are estimated from cancer cases recorded by the Belgian Cancer Registry. The obtained results suggest that a new offer could develop, targeting the particular needs of cancer patients.

**Keywords:** Critical illness, medical insurance, right to be forgotten, multistate models.

## 3.1 Introduction and motivation

Massart (2018) testimonial illustrates difficulties faced by cancer survivors to access life and health insurance products. See also Hendriks et al. (2021) for the particular case of childhood cancer survivors. However, progress in medicine over the last 20 years greatly improved the prognosis of several types of cancer. In parallel, many tools have been developed in biostatistics and epidemiology to study mortality and morbidity associated with cancer. These advances have already resulted in a better access to insurance products for people diagnosed with cancer. For example, this led France and Belgium to establish a “*droit à l’oubli*” (translated literally as “right to be forgotten” in the remainder of this text) granting access to insurance after a waiting period of at most 10 years starting at the end of the successful therapeutic protocol, in absence of relapse within this period. The 10-year period has even been shortened for several types of cancer with a good prognosis. We refer the reader to Soetewey et al. (2021) for an actuarial analysis of the “right to be forgotten” mechanism. Several initiatives purpose to extend this right throughout EU countries, by law or through a convention with insurance sector (as in Luxembourg). We refer the reader e.g. to Scocca and Meunier (2020) as well as to “Survivorship challenge 3.4: Lack of knowledge of the stigma associated with cancer” listed in Lawler et al. (2021) purposing to take

advantage of the existing legal framework in four EU member countries (France, Belgium, Luxembourg and the Netherlands) to investigate a pan-European legal framework on access to financial services for cancer survivors.

Although the establishment of such a “right to be forgotten” is a clear progress for cancer survivors, there is still room for improvement by filling the coverage gap during the waiting period opening this right. Social security and supplemental health insurance cover most medical costs related to cancer, but non-medical costs are usually paid out of pocket. The latter include lost income (the part not covered by Social security or supplemental disability insurance), travel to and from hospital (especially for patients living in rural areas), travel and family lodging expenses, deductibles and co-payments, non-conventional comfort treatments, private nursing care costs, and non-nursing help with activities of daily living. They can rapidly become a financial hardship, even for patients who are treated on an outpatient basis.

Some insurance companies market supplemental cancer insurance policies that pay lump sum benefits upon diagnosis of cancer or a temporary life annuity to face these out-of-pocket expenditures. These products developed in Asia, North America and UK (Bennett et al., 1998; Nielsen and Mayer, 2000). Using Taiwan National Health Insurance Database, Yue et al. (2018) priced two types of whole-life insurance products: (i) a (lump-sum) benefit paid when the insured is diagnosed with cancer for the first time (and the contract terminates after the benefit is paid) and (ii) an annual benefit paid after the insured is diagnosed with cancer and as long as he or she survives. Shang (2019) considered term life products and evaluated the extra cost related to cancer. Let us also mention that cancer is also typically included into the diseases covered by critical illness insurance policies.

The products considered in this paper are specifically related to the waiting period opening the “right to be forgotten”, with temporary covers restricted to that period to fill the gap in coverage on a market where such a right has been implemented. First, stand-alone products are studied, including cancer insurance with lump sum payment at diagnosis, or temporary life annuity starting at diagnosis. In the latter case, periodic payments may correspond to insurance premiums of another product, or even to loan reimbursement. Then, riders included in a package are discussed. We consider term insurance with accelerated death benefit paid as a lump sum at diagnosis or as a temporary life annuity starting at diagnosis. The payment of rider reduces death benefit specified in term insurance. Finally, we discuss products granting access to some specific insurance cover (such as mortgage insurance) during the waiting period opening the “right to be forgotten”. This is especially important at young age, to guarantee access to property and home ownership (in case of house loan) and to entrepreneurship (in case of professional loan) to cancer patients whose health status has improved but who cannot benefit from the “right to be forgotten” because the waiting period is not exhausted. This guarantee can be bought by parents for their children (as it is commonly the case in medical insurance, where extra premiums ensure that children can continue their parents’ medical cover when they leave the household, whatever their future health status) or by young adults starting their professional career (in supplement to individual health insurance, for instance). The cover may be subject to some deferred period after diagnosis in order to lower its cost (without real impact since it is very unlikely that cancer patients consider buying a house or developing professional activities right after diagnosis). The price of this cover also allows to evaluate the cost of offering standard premium rates to all applicants in mortgage insurance related to property loans. This is the approach followed by Crédit Mutuel in France, which announced in November 2021 the end of health questionnaire when applying for a home loan (under some conditions, like age below 62, amount borrowed less than 500,000 euros, having Crédit Mutuel as main bank since 7 years at least, this duration serving as an implicit waiting period

as in the “right to be forgotten” mechanism, and electing domicile in the house to be bought). The bank even announces that this decision will have a retroactive effect for the ongoing loans. The loss in mortgage insurance premiums is evaluated to 70 millions euros per year for Crédit Mutuel. Our calculation help to assess the cost of this decision.

Cancer typically belongs to the set of critical illnesses and is thus also covered under several standard insurance policies. Long-term care or disability insurance policies also pay benefits in case of cancer. However, depending on the terms of the contract, a cancer patient may well have completed his or her treatment before the deferred period required to receive insurance benefits has been reached. Some patients even continue working while receiving chemotherapy or radiation therapy and thus do not qualify for loss of income. The cover comprised in the products considered in this paper is activated at diagnosis. These policies are targeted to help patients facing out-of-pocket expenditures while treatment occurs or to grant them access to other insurance products during the waiting period opening the “right to be forgotten”. They thus constitute a new offer on a market where such a right has been implemented (as it is the case in Belgium or France, and may be generalized at EU level).

To illustrate our proposals, we consider the Belgian market where the “right to be forgotten” has been inserted in the Insurance Law in April 2019, with reduced waiting periods for some cancer types defined by Royal Decree in May 2019. Calculations are based on data available from the Belgian Cancer Registry (BCR), a national population-based cancer registry collecting data on all new cancer diagnoses in Belgium since the incidence year 2004. The paper considers three cancers with clear differences in terms of incidence, survival after diagnosis, and waiting periods defined by Royal Decree: melanoma (ICD-10 C43) with waiting period reduced up to 1 year after the end of the successful therapeutic protocol, thyroid (ICD-10 C73) with waiting period reduced up to 3 to 6 years after the end of the successful therapeutic protocol, and female breast (ICD-10 C50) cancer subject to the standard 10-year waiting period. Melanoma and thyroid cancer patients are known to have limited excess mortality compared to the general population (Soetewey et al., 2021). This has been recognized by reducing the waiting period opening the “right to be forgotten” in the Royal decree published in the Belgian Official Journal on June 14, 2019. The situation for female breast cancer patients is different with usually a high survival probability in the first years after diagnosis before it eventually decreases due to late cancer recurrences.

In this paper we perform our actuarial calculations in a 3-state Semi-Markov model assuming that a policyholder can be either “healthy”, “ill” (diagnosed with cancer), or “dead”. For the sake of easiness, only transitions from healthy to ill, healthy to dead and ill to dead are allowed so that cancer is assumed to be permanent (i.e., no recovery is possible). This non-reversibility greatly simplifies the computations (as the 3-state process is hierarchical) and appears to be reasonable for cancer insurance considered in this paper, with temporary cover restricted to the waiting period opening the “right to be forgotten” and benefits paid after the first diagnosis. For premium calculations, if mortality in cancer state reverts to the standard level after a sufficiently long period (like in cure models) then this is equivalent to a transition back to the initial healthy state. This model has been considered by Denuit et al. (2019) for long-term care insurance. Let us mention that Dębicka et al. (2015) also considered multistate models to combine reverse annuity contracts with critical illness (in fact cancer) insurance for retired people. The cancer state is splitted into several stages to capture duration effects while remaining with a Markov structure. Here, we consider younger ages and focus on the waiting period opening the “right to be forgotten”, performing calculations in the 3-state Semi-Markov model.

The remainder of the paper is organized as follows. Section 2 describes the 3-state



Figure 3.1: Semi-Markov 3-state model for cancer insurance.

Semi-Markov model used for premium calculation. Cancer insurance products are described in Section 3, where corresponding premium rates are computed in the model of Section 2. The final Section 4 discusses the results and concludes.

## 3.2 Semi-Markov 3-state model

### 3.2.1 State space and transitions

Multistate models offer a convenient representation for life and health insurance liabilities when benefits are associated to sojourns in, or transitions between different states (Dickson et al., 2013; Pitacco, 2014). We consider an individual aged  $x$  at policy issue, taken as time 0. His or her history is described by the stochastic process  $\{X_t, t \geq 0\}$  where  $X_t$  gives the state occupied at time  $t$ . Here,  $t$  corresponds to contract seniority. In this paper, we consider that  $X_t \in \{a, i, d\}$  where state  $a$  stands for “active” (healthy), state  $i$  stands for “ill” (cancer) and state  $d$  stands for “dead” as represented in Figure 3.1. At the time of diagnosis, individual moves from state  $a$  to state  $i$ . We do not allow for recovery (but mortality rates in state  $i$  may ultimately become similar to those applying in state  $a$  for cured individuals). Since benefits only relate to the first diagnosis, state  $i$  corresponds here to the first time a policyholder is diagnosed with cancer.

The time spent in state  $i$  is known to influence mortality so that we introduce the random variable  $Z_t$ , defined as the time spent in the state occupied at time  $t$ . Formally,

$$Z_t = \max\{z \leq t \mid X_t = X_{t-h} \text{ for all } 0 \leq h \leq z\}.$$

For an individual in state  $i$  at time  $t$ ,  $Z_t$  is the time since diagnosis. Henceforth, we work under the Semi-Markov assumption: only the current state  $X_t$  and the time  $Z_t$  spent in the current state influence future transitions. This means that stochastic process  $\{(X_t, Z_t), t \geq 0\}$  is a Markov process.

### 3.2.2 Transition intensities

Transition intensities quantify the instantaneous risk of making a given transition, depending on the state currently occupied and sojourn time. Assuming that  $Z_t$  only matters in state  $i$ , transition intensities are defined by the following limits:

$$\begin{aligned} \mu_{x+t}^{ai} &= \lim_{h \rightarrow 0} \frac{P[X_{t+h} = i \mid X_t = a]}{h} \\ \mu_{x+t}^{ad} &= \lim_{h \rightarrow 0} \frac{P[X_{t+h} = d \mid X_t = a]}{h} \\ \mu_{x+t;z}^{id} &= \lim_{h \rightarrow 0} \frac{P[X_{t+h} = d \mid X_t = i, Z_t = z]}{h}, z < t. \end{aligned}$$

State  $a$  remains Markovian so that transition intensities from that state do not depend on the time spent in the state, but only on attained age  $x + t$ . On the contrary,

there is an influence of the duration of stay in state  $i$  so that transition intensities from state  $i$  depend on both attained age  $x + t$  and time  $z$  since diagnosis.

### 3.2.3 Data

#### 3.2.3.1 Belgian Cancer Registry (BCR)

We consider the data available from the Belgian Cancer Registry (BCR), a national population-based cancer registry collecting data on all new cancer diagnoses in Belgium since the incidence year 2004. For the execution of this main task, the BCR relies on its own specific legislation (more information can be found on the BCR website, at [kankerregister.org](http://kankerregister.org)).

To illustrate our work, we restrict our analyses to three cancer types: melanoma (ICD-10 C43), thyroid (ICD-10 C73) and female breast (ICD-10 C50) cancer. These three cancer sites have been selected to evaluate the proposed insurance products in different scenarios. Melanoma and thyroid cancer patients are known to have a limited excess mortality compared to the general population. The situation for female breast cancer patients is different with usually a high survival probability in the first years after the date of diagnosis before it eventually decreases due to late cancer recurrences. We consider only female breast cancer as there are very few registrations regarding to male breast cancer. We also limit our analyses to patients aged 20 to 69 at diagnosis since the products considered in this paper target young adults and active life.

A total of 24,325 persons were diagnosed with melanoma, 10,789 with thyroid and 105,127 with breast cancer between 2004 and 2018, and were followed-up until April 1, 2020. Follow-up thus ranged from 2 to 16 years. Only one record per patient (with the earliest incidence date) within each cancer site was kept for patients with multiple primary diagnoses. This is in accordance with the insurance products under consideration which are activated at the time of the first diagnosis. A minority of patients without national security number were excluded from the analysis. Patients lost to follow-up (mostly due to moving abroad) and patients still alive at the end of the follow-up period were treated as censored observations.

Table 3.1 summarizes the number of included cases, number and proportion of deaths and percentage of lost to follow-up before April 1, 2020 per type of cancer, gender and age group. The fraction of patients lost to follow-up per subgroup varied from 1.06% for women with melanoma cancer aged 35-49 to 4.04% for male thyroid cancer patients aged 20-34. The total fraction of patients lost to follow-up cases, regardless of gender, site or age group was 1.4%.

#### 3.2.3.2 General population

The products considered in this paper are sold to individuals before diagnosis (thus, in state  $a$ ). This is in contrast with the study by Soetewey et al. (2021) which considered individuals in state  $i$ . Therefore, we also need mortality in the general population. Belgian population life tables are available from Statbel (the Belgian statistical office) and can be freely downloaded from the website [statbel.fgov.be](http://statbel.fgov.be).

### 3.2.4 Estimation

Transition intensities are often assumed to be piecewise constant in order to ease actuarial calculations. Starting from state  $a$ , this means that the identities

$$\mu_{x+k+t}^{ai} = \mu_{x+k}^{ai} \text{ and } \mu_{x+k+t}^{ad} = \mu_{x+k}^{ad} \quad (3.1)$$

Gender	Cancer site	Age at diagnosis	Lost to follow-up	Number of included cases	Number of deaths
Men	Melanoma	20-34	3.38%	857	83
		35-49	2.24%	2,812	354
		50-69	1.70%	6,000	1,268
	Thyroid	20-34	4.04%	322	6
		35-49	3.21%	841	58
		50-69	1.92%	1,563	297
Women	Melanoma	20-34	3.22%	2,174	70
		35-49	1.06%	5,267	317
		50-69	1.12%	7,215	874
	Thyroid	20-34	3.76%	1,435	10
		35-49	2.51%	3,029	78
		50-69	1.83%	3,599	390
	Breast	20-34	2.87%	2,685	423
		35-49	1.49%	29,007	3,419
		50-69	1.07%	73,435	12,730
Total			1.40%	140,241	20,377

**Table 3.1: Number of persons diagnosed with melanoma, thyroid and female breast cancer in Belgium between 2004 and 2018 (BCR data) by gender, site and age group, together with the percentage of lost to follow-up and the number of deaths.**

hold for every integers  $x$  and  $k$  and fractional  $0 \leq t < 1$ . Transition intensity from state  $i$  is displayed as a function depending on the age at diagnosis and the time elapsed since diagnosis, i.e.

$$\mu_{x+\xi;z}^{id} = \tilde{\mu}(x + \lfloor \xi - z \rfloor, \lfloor z \rfloor) \quad (3.2)$$

for some given function  $\tilde{\mu}$  with integer arguments, where  $\lfloor \cdot \rfloor$  denotes rounding from below (i.e., the integer part). The arguments of  $\tilde{\mu}(\cdot, \cdot)$  represent, respectively, integer parts of age at entry in the ill state and time spent in that state. We thus work with age last birthday at diagnosis and the number of years since diagnosis, rounded from below. Of course, more accurate calculations can be performed by refining the time step, if needed.

When intensities are piecewise constant, they are easily estimated by the ratio of the observed number of transitions (diagnosis or death) to the corresponding exposure (in the state to be left). Precisely, consider a given integer age  $y$  and let  $N_y^{ai}$  be the number of transitions from state  $a$  to state  $i$ , that is, the number of diagnoses, among individuals aged  $y$  last birthday. Similarly, let  $N_y^{ad}$  be the number of transitions from state  $a$  to state  $d$ , that is, the number of deaths recorded among healthy individuals aged  $y$  last birthday. Let  $E_y^a$  denote the (central) exposure to risk in state  $a$ , that is, the time spent by all individuals aged  $y$  last birthday in state  $a$ . Because general population mortality statistics do not record exposures but only the number of individuals at the beginning of the period (or initial exposure to risk), we assume that transitions occur in the middle of the period. Under (3.1), the maximum likelihood estimators of  $\mu_y^{ai}$  and  $\mu_y^{ad}$  are respectively given by  $N_y^{ai}/E_y^a$  and  $N_y^{ad}/E_y^a$  and these values apply between ages  $y$  and  $y + 1$ . Similar formulas hold true under (3.2) for estimating  $\tilde{\mu}(y, z)$  for integer values  $y$  and  $z$ , classifying transitions and recording exposures according to age last birthday  $y$  and integer number  $z$  of years since diagnosis.

Given that BCR covers the whole population, it would be possible to subtract from exposures and death counts available from Statbel the time lived and the number of



deaths among cancer patients. In this way, the estimated  $\mu_y^{ad}$  would only account for mortality not related to the cancer under consideration. However, in this paper, estimated intensities  $\mu_y^{ad}$  have been obtained from the Belgian population life tables, so ignoring the influence of cancer mortality. The fact that population life tables include cancer mortality is not an issue as mortality for a given cancer represents only a small fraction of the overall mortality, and correcting for this over-representation of the cancer being studied has, in practice, an insignificant effect. The transition intensities from healthy to ill (cancer) and from ill to death can be estimated from BCR data (combined with general population exposures in the former case).

Even if the general shape of the mortality and incidence curves is generally clearly visible, erratic variations often remain. As long as these random departures do not reveal anything about the underlying mortality or morbidity pattern, they should be removed before entering actuarial calculations. This process is known as graduation in the actuarial literature. As there is no simple parametric model able to capture the structure of mortality and morbidity, actuaries generally use a generalized additive regression model with Poisson response distribution for transition counts, see e.g. Denuit and Legrand (2018). The method can be summarized as follows. Under (3.1), maximum likelihood inference can be equivalently conducted under the hypothesis that  $N_y^{ai}$  is Poisson distributed with mean  $E_y^a \mu_y^{ai}$ . The transition intensity  $\mu_y^{ai}$  is then represented as  $\ln \mu_y^{ai} = s(y)$  for some smooth function  $s(\cdot)$  to be estimated from the data, under the assumption that  $N_y^{ai}$ ,  $y = 20, 21, \dots, 69$  are mutually independent. The function  $s(\cdot)$  is estimated with the help of a Poisson generalized additive model and the resulting estimate is used to produce the transition rates adopted to perform all actuarial calculations in the remainder of this paper. A similar procedure is followed to produce the other transition intensities entering the calculations. The resulting estimated transition intensities are visible in Figures 3.2-3.4.

Figure 3.2 displays the estimated intensities  $\mu_y^{ad}$  as functions of attained age  $y$  for males and females. Belgian regulatory life tables XR and XK are also displayed there: life table XK defines minimum premium amount for life insurance policies with a positive sum at risk (thus comprising mainly death benefits) whereas life table XR defines minimum premium amount for policies with a negative sum at risk (thus comprising mainly survival benefits). Life table XK is conservative and generates a relatively high safety loading. Dating back to the 1990s, life table XR does not comprise any safety loading anymore for women (but since it only defines minimal premium amounts, insurers remain free to charge higher premiums to remain solvent). We recognize on Figure 3.2 the exponential increase in mortality at adult ages (the accident hump is not visible because actual values are displayed along the vertical axis, without log transform).

Figure 3.3 displays the estimated intensities  $\mu_y^{ai}$  as functions of attained age  $y$  for the three types of cancer considered in this paper, separately for males and females. We can see there that incidence curves greatly differ among the three cancer types under consideration. In particular, after age 30, incidence for women breast cancer largely exceeds the one for melanoma and thyroid. For males, incidence rates are closer at young ages but exhibit different age trends.

Estimated  $\tilde{\mu}(y, z)$  (with  $y$  and  $z$  corresponding to integer part of age at diagnosis and time since diagnosis, respectively) is displayed in Figure 3.4. We can see there that mortality increases with age at entry and sojourn time for both genders and all three considered cancer sites. Mortality, however, increases less rapidly with sojourn time for young patients compared to old patients. We also see that, for patients below age 40, mortality remains low even after a long period after diagnosis (i.e., for large values of sojourn time). Note that since we computed these quantities considering all causes of death, they account for both mortality from the cancer of interest and mortality from other causes. This is in line with the application to insurance since



Figure 3.2: Estimated transition intensities  $\mu_y^{ad}$  as functions of attained age  $y$ . General population (Statbel, continuous line) and insurance regulatory life tables XR (broken line) and XK (dotted line).



Figure 3.3: Estimated transition intensities  $\mu_y^{ai}$  as functions of attained age  $y$ , for different cancer types.

benefits do not vary according to the cause of death for the products considered in this paper.

### 3.2.5 Transition probabilities

The following probabilities are useful to perform actuarial calculations. Considering an individual who is healthy at age  $x + t$ , that is, who is in state  $a$  at time  $t$ , the probability of being in state  $i$  at time  $t + u$  is denoted as

$${}_u p_{x+t}^{ai} = P[X_{t+u} = i | X_t = a],$$

the probability of being in state  $d$  at time  $t + u$  is denoted as

$${}_u p_{x+t}^{ad} = P[X_{t+u} = d | X_t = a],$$

and the probability of being in state  $a$  at time  $t + u$  is denoted as

$${}_u p_{x+t}^{aa} = P[X_{t+u} = a | X_t = a].$$

Since the time spent in state  $i$  influences future transitions, the random variable  $Z_t$  also enters the transition probabilities from that state. Precisely, considering an ill individual aged  $x + t$  who has been diagnosed at time  $t - z$ , that is, who is in state  $i$  at time  $t$  since time  $t - z$ , the probability of being in state  $d$  at time  $t + u$  is denoted as

$${}_u p_{x+t;z}^{id} = P[X_{t+u} = d | X_t = i, Z_t = z]$$

and the probability of being in state  $i$  at time  $t + u$  is denoted as

$${}_u p_{x+t;z}^{ii} = P[X_{t+u} = i | X_t = i, Z_t = z].$$

By assumption, recovery is not possible. Hence, transition probabilities  ${}_u p_{x+t}^{aa}$  and  ${}_u p_{x+t;z}^{ii}$  are in reality sojourn probabilities, i.e.

$$\begin{aligned} {}_u p_{x+t}^{aa} &= P[X_{t+h} = a \text{ for all } 0 < h \leq u | X_t = a] \\ {}_u p_{x+t;z}^{ii} &= P[X_{t+h} = i \text{ for all } 0 < h \leq u | X_t = i, Z_t = z]. \end{aligned}$$

Sojourn probabilities are easy to compute when transition intensities are piecewise constant, as exponential functions of minus integrated exit rates (or cumulative hazards). This property will be used repeatedly in the next section.

## 3.3 Cancer insurance products

### 3.3.1 Notation and specific policy conditions

Henceforth,  $v(s, t)$  is the present value at time  $s$  of a unit payment made at time  $t$  (with  $s < t$  and  $v(s, s) = 1$ ). We assume that the technical interest rate used in actuarial calculation is constant and we denote as  $\delta$  the corresponding instantaneous force of interest, that is,  $v(s, t) = \exp(-\delta(t - s))$ . Also, we denote as  $\mu_{x+t}^{a\bullet}$  the exit intensity from state  $a$ , that is,  $\mu_{x+t}^{a\bullet} = \mu_{x+t}^{ad} + \mu_{x+t}^{ai}$ .

In this paper, we consider temporary covers where diagnosis has to occur within the next  $n$  years to get the insurance benefit. The insured period is thus the time interval  $[0, n]$ , in the sense that a benefit is payable only if the time of diagnosis belongs to this interval. In principle, the insured period begins at policy issue and ends at policy termination, subject to the following specific policy conditions. The waiting period (or “elimination” period)  $w$  is the period following the policy issue during which the insurance cover is not yet operating. The waiting period aims at limiting



(a) Women with melanoma cancer



(b) Men with melanoma cancer



(c) Women with thyroid cancer



(d) Men with thyroid cancer



(e) Women with breast cancer

Figure 3.4: Mortality intensities according to age at diagnosis and sojourn time in the cancer state.

the effects of adverse selection, in particular because of pre-existing insured's health conditions. The waiting period considered here is the one specified in the contract, not to be confused with the waiting period opening the "right to be forgotten" fixed by the law. Sometimes, the benefit is not payable at diagnosis but only if the policyholder has survived a certain minimum period after diagnosis, called the deferred period and denoted as  $f$ . This policy condition acts as a deductible and essentially purposes to reduce the cost and hence the premium of the insurance product; premium reduction can be particularly significant in case high mortality immediately follows diagnosis.

All premium calculations are performed on a unisex basis, in accordance with EU regulation which prohibits any difference in insurance cover or amount of premium by gender. Transitions observed for males and females have therefore been combined to produce a set of intensities independent of gender, which are used for all calculations performed in this section. For computations with respect to breast cancer, transitions for women only have been considered as only female breast cancer cases are included.

Let us comment on the particular case of the breast cancer cover. Even if this contract targets female policyholders, male breast cancer, though rare, does exist. To be consistent with the anti-discrimination EU directive, the coverage has to be offered to both males and females and priced under unisex basis. Notice that a unisex tariff may well be based on women's data exclusively (the requirement is that the premium cannot differ between male and female policyholders). Even if breast cancers are generally rare for males, they often have a very poor prognosis. This is because (i) these cancers are often detected at a later stage compared to women, (ii) there is not much research devoted to breast cancer affecting males and (iii) available treatments against female breast cancer are difficult to adapt to treat male patients because of the marked difference in hormonal status. According to the BCR, the 5-year prevalence from 2013 to 2017 in Belgium is 398 for men compared to 47,423 for women.

Since this cover may raise some concerns related to gender-based discrimination, let us consider the Femina cover sold in Belgium by AG Insurance (one of the market leaders). This is a sickness insurance product with lump sum benefits paid after diagnosis of some specific cancers affecting women, including breast cancer that may also affect men. As its name indicates, this product clearly targets women. The policy must be issued before the age of 60, offering a lifelong cover (subject to a severe underwriting conditions). Health Minister had to answer some specific queries by members of the Belgian Parliament about possible discriminatory issues related to Femina (see Question 7 by Deputy Karin Jiroflée to Minister Kris Peeters about Femina insurance, ref. P2321, plenary session of October 5, 2017). Minister Peeters asked the Financial Services and Markets Authority (FSMA, protecting consumers in the financial sector, including bank and insurance) to determine whether this product complies with the anti-discrimination EU Directive. The conclusion of the FSMA study was provided during the meeting of Parliament Economy Commission of March 28, 2018. Since AG Insurance indicated that Femina cover could also be bought by men, no discriminatory issue was found in relation to the product. This shows that even if products are marketed to address specific needs of the male or female population and are priced accordingly using data from the targeted population, this does not violate the EU Directive as long as both genders can access the cover at the same conditions (even if it is very unlikely that males will ever buy the Femina insurance cover).

#### 3.3.2 Stand-alone covers

##### 3.3.2.1 Lump sum

A first possibility is to pay a lump sum at diagnosis. The beneficiary can use this amount to face out-of-pocket expenditures related to treatment. The expected present

value of a unit lump sum paid at diagnosis is given by

$$\bar{A}_{x;n]^{a;a \rightarrow i}} = \int_0^n v(0, t) {}_t p_x^{aa} \mu_{x+t}^{ai} dt.$$

In case the contract specifies a waiting period  $w$ , the integral is over  $(w, n)$  instead of  $(0, n)$ . Since diagnosis is recorded in the BCR database, the payment date  $t$  is easy to check.

When transition intensities are piecewise constant, we get

$$\begin{aligned} \bar{A}_{x;n]^{a;a \rightarrow i}} &= \mu_x^{ai} \frac{1 - \exp(-\delta - \mu_x^{a\bullet})}{\delta + \mu_x^{a\bullet}} \\ &+ \sum_{j=1}^{n-1} \mu_{x+j}^{ai} \exp\left(-\sum_{k=0}^{j-1} \mu_{x+k}^{a\bullet} - j\delta\right) \frac{1 - \exp(-\delta - \mu_{x+j}^{a\bullet})}{\mu_{x+j}^{a\bullet} + \delta}. \end{aligned} \quad (3.3)$$

In principle, the payment could be deferred. In many policies, the benefit is not payable until the need has lasted a certain minimum period called the deferred period (Pitacco, 2014). Here, this would mean that the lump sum is not paid at diagnosis but the payment is deferred later on. This may not be desired by the customers buying the product considered here so that we do not consider this possibility. Deferred periods may nevertheless be useful to lower premiums in case they are too expensive.

The values of  $\bar{A}_{x;n]^{a;a \rightarrow i}}$  obtained in the Semi-Markov 3-state model are displayed in Figure 3.5 for ages  $x \in \{20, 21, \dots, 40\}$ , coverage period  $n = 20$  years, and yearly interest rate 1%, that is,  $\delta = \ln 1.01$ . Without discounting, that is, setting  $\delta = 0$  or  $v(s, t) = 1$  for all  $s < t$ ,  $\bar{A}_{x;n]^{a;a \rightarrow i}}$  is the probability of being diagnosed with cancer for a healthy individual aged  $x$  over the next  $n$  years. Remember that for melanoma and thyroid cancers, it should be interpreted as the probability on the whole population while for breast cancer, it should be interpreted as the probability only among women. We can see on Figure 3.5 that premium amounts remain rather low for melanoma and thyroid cancers, but considerably increase for breast cancer because of larger incidence within the Belgian population (culminating at 0.017 per unit of sum insured at age 40).

**Remark 3.3.1.** The single premium  $\bar{A}_{x;n]^{a;a \rightarrow i}}$  can be converted into a periodic one by dividing it with

$$\bar{a}_{x;n]^{aa}} = \int_0^n v(0, t) {}_t p_x^{aa} dt, \quad (3.4)$$

with the understanding that the premium is payable until diagnosis or death.

### 3.3.2.2 Temporary life annuities

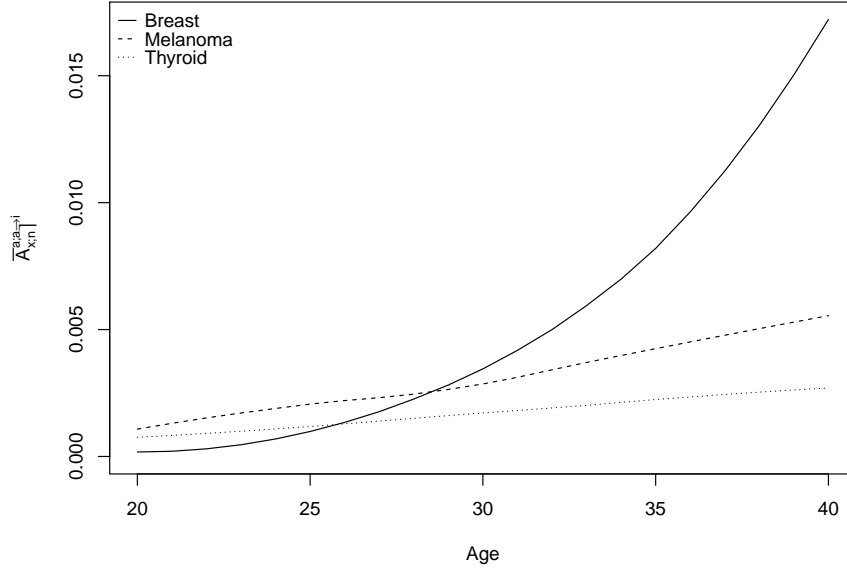
Insured benefits can also consist in a temporary life annuity starting at diagnosis. This provides cancer patient with a periodic income to face out-of-pocket expenditures. The corresponding expected present value if payments are made continuously at a constant unit rate as long as cancer patient survives is given by

$$\bar{a}_{x;n]^{ai}} = \int_0^n {}_t p_x^{aa} \mu_{x+t}^{ai} v(0, t) \bar{a}_{x+t;0}^{ii} dt \quad (3.5)$$

where

$$\bar{a}_{x+t;0}^{ii} = \int_0^m {}_s p_{x+t;0}^{ii} v(t, t+s) ds \quad (3.6)$$

with  $m$  denoting the maximal payment duration. Here,  $m$  is given in policy conditions and may vary with the waiting period opening the “right to be forgotten” by the law, that is, it may depend on cancer type.



**Figure 3.5: Values of  $\bar{A}_{x:n}^{a;a \rightarrow i}$  as function of age  $x \in \{20, 21, \dots, 40\}$  for different cancer types with  $n = 20$  and yearly interest rate 1%.**

The duration  $m$  could be determined in two ways, at least. Either we adopt the reduced waiting periods specified by Royal decree but it would then be necessary to add the duration of treatment since the “right to be forgotten” in the law starts at the end of a successful treatment protocol. Or, we take the duration of the modified “right to be forgotten” since diagnosis as determined by Soetewey et al. (2021). While it could have been interesting to formally compare both approaches, individual data on the type and length of treatment for each case is not reported in the BCR and such information is not readily available. Even if it were available, the definition of the end of the treatment remains unclear (and this is precisely the reason why Soetewey et al. (2021) suggested to let the waiting period start from diagnosis, to avoid endless disputes when a claim occurs). Moreover, durations of treatment are heterogeneous even within the same cancer type, usually unpredictable. Optimal durations are often still open to debates, see e.g. Schvartsman et al. (2019), making it hard to include the duration of treatment in the actuarial computations. In any case, a reduction in treatment length due to the progress made in medical treatment of cancer would obviously lead to closer agreement between the two approaches. Since the date of diagnosis, as recorded in national registries, offers the great advantage of not being subject to any discussion and to allow the patient to know from the start when the waiting period will end, we think that all parties benefit from using the date of diagnosis instead of the end of treatment. For these reasons, we favor the second approach in the present paper.

When transition intensities are piecewise constant, we get

$$\begin{aligned}\bar{a}_{x;n}^{ai} &= \sum_{j=0}^{n-1} \int_j^{j+1} {}_t p_x^{aa} \mu_{x+t}^{ai} v(0, t) \bar{a}_{x+t;0}^{ii} dt \\ &= \sum_{j=0}^{n-1} {}_j p_x^{aa} v(0, j) \int_0^1 {}_t p_{x+j}^{aa} \mu_{x+j+t}^{ai} v(j, j+t) \bar{a}_{x+j+t;0}^{ii} dt\end{aligned}\quad (3.7)$$

for  $j \in \{0, \dots, n-1\}$  and  $t \in [0, 1)$ , where

$$\begin{aligned}\bar{a}_{x+j+t;0}^{ii} &= \int_0^m {}_s p_{x+j+t;0}^{ii} v(j+t, j+t+s) ds \\ &= \sum_{k=0}^{m-1} \int_k^{k+1} {}_s p_{x+j+t;0}^{ii} v(j+t, j+t+s) ds \\ &= \sum_{k=0}^{m-1} {}_k p_{x+j+t;0}^{ii} v(j+t, j+t+k) \\ &\quad \int_0^1 {}_s p_{x+j+t+k;k}^{ii} v(j+t+k, j+t+k+s) ds \\ &= \sum_{k=0}^{m-1} \exp\left(-\sum_{l=0}^{k-1} \tilde{\mu}(x+j, l) - k\delta\right) \int_0^1 \exp(-s(\tilde{\mu}(x+j, k) + \delta)) ds \\ &= \frac{1 - \exp(-\delta - \tilde{\mu}(x+j, 0))}{\delta + \tilde{\mu}(x+j, 0)} + \sum_{k=1}^{m-1} \exp\left(-\sum_{l=0}^{k-1} \tilde{\mu}(x+j, l) - k\delta\right) \\ &\quad \frac{1 - \exp(-\delta - \tilde{\mu}(x+j, k))}{\delta + \tilde{\mu}(x+j, k)} \\ &= \bar{a}_{x+j;0}^{ii}\end{aligned}\quad (3.8)$$

which shows that  $\bar{a}_{x+j+t;0}^{ii}$  can be taken out of the integral in the expression of  $\bar{a}_{x;n}^{ai}$ . Hence, we have

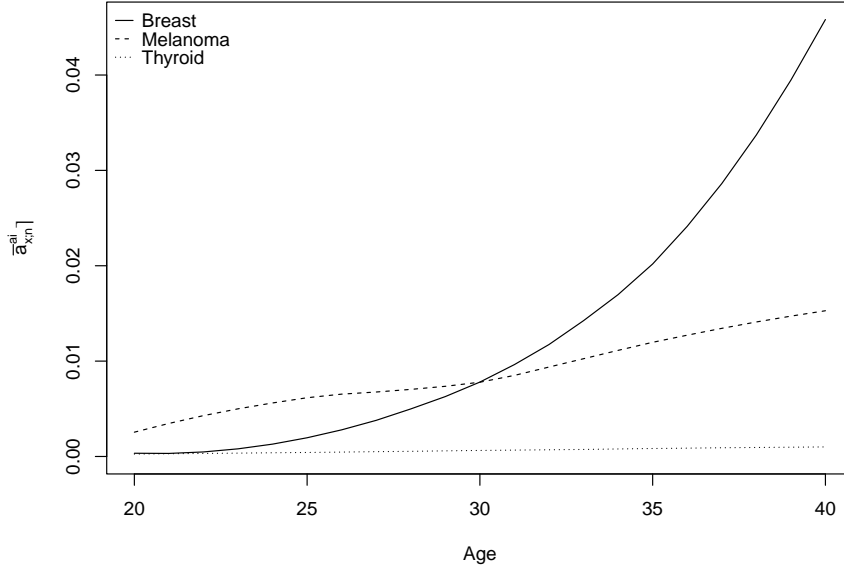
$$\begin{aligned}\bar{a}_{x;n}^{ai} &= \sum_{j=0}^{n-1} {}_j p_x^{aa} v(0, j) \bar{a}_{x+j;0}^{ii} \int_0^1 {}_t p_{x+j}^{aa} \mu_{x+j+t}^{ai} \exp(-t\delta) dt \\ &= \mu_x^{ai} \bar{a}_{x;0}^{ii} \frac{1 - \exp(-\delta - \mu_x^{a\bullet})}{\delta + \mu_x^{a\bullet}} \\ &\quad + \sum_{j=1}^{n-1} \mu_{x+j}^{ai} \bar{a}_{x+j;0}^{ii} \exp\left(-\sum_{k=0}^{j-1} \mu_{x+k}^{a\bullet} - j\delta\right) \frac{1 - \exp(-\delta - \mu_{x+j}^{a\bullet})}{\delta + \mu_{x+j}^{a\bullet}}.\end{aligned}\quad (3.9)$$

Values of  $\bar{a}_{x;n}^{ai}$  are computed as a function of age at policy issue  $x \in \{20, 21, \dots, 40\}$  with  $n = 20$  and a yearly technical interest rate of 1%, that is,  $\delta = \ln 1.01$ . The duration  $m$  is taken to be equal to 9 years for melanoma and 1 year for thyroid, in accordance with the reduced waiting periods fixed in the Belgian law for these cancers. For breast cancers,  $m$  is taken to be the standard waiting period of 10 years. The numerical values are displayed in Figure 3.6. We can see there that the product is cheaper for melanoma and thyroid cancers from age 30, thanks to lower incidence rates and payment duration. Moreover, premiums increasing more rapidly with age are obtained for breast cancer, because of higher incidence rates and payment duration.

### 3.3.3 Combined products

Combined products correspond to insurance packages where cancer insurance supplements a reference cover. Several examples are described hereafter.





**Figure 3.6:** Values of  $\bar{a}_{x;n}^{ai}$  as function of age  $x$  for different cancer types with  $n = 20$  and yearly interest rate 1%.

### 3.3.3.1 Premium exemption

A first possibility is to use the temporary life annuity starting at diagnosis for paying the premiums of a reference cover, or even to reimburse a loan secured by mortgage insurance, for instance. Calculations are performed as explained before, with benefits matching amounts of premium or loan reimbursement.

### 3.3.3.2 Term-life insurance with cancer acceleration benefit

Cancer insurance can be added as a rider to a term-life insurance policy. In this case, the amount of death benefit is (totally or partially) converted into a lump sum paid at diagnosis. Specifically, let

$$\bar{A}_{x;n}^{a;a \rightarrow d} = \int_0^n v(0, t) {}_t p_x^{aa} \mu_{x+t}^{ad} dt \quad (3.10)$$

be the expected present value of a unit lump sum paid at death occurring in state  $a$ . In practice, this amount is replaced with the XK premium if the latter is higher. Let  $c_{ad}$  be the amount of death benefit for a policyholder in state  $a$ . A proportion  $\alpha \in [0, 1]$  of the death benefit may be paid at diagnosis and the remaining  $1 - \alpha$  at death in state  $i$ . For a unit death benefit, the expected present value of insurance benefits is then given by

$$\begin{aligned} \bar{A}_{x;n}^{(\alpha)} &= \bar{A}_{x;n}^{a;a \rightarrow d} \\ &+ \int_0^n {}_t p_x^{aa} \mu_{x+t}^{ai} \left( \alpha v(0, t) + \int_0^{n-t} {}_z p_{x+t;0}^{ii} \mu_{x+t+z}^{id} (1 - \alpha) v(0, t + z) dz \right) dt. \end{aligned} \quad (3.11)$$

When transition intensities are piecewise constant, we can compute the pure premium as follows. First, the part of the pure premium corresponding to death

benefits for a healthy individual writes

$$\begin{aligned} \bar{A}_{x;n] }^{a;a \rightarrow d} &= \mu_x^{ad} \frac{1 - \exp(-\delta - \mu_x^{a\bullet})}{\delta + \mu_x^{a\bullet}} \\ &\quad + \sum_{j=1}^{n-1} \mu_{x+j}^{ad} \exp\left(-\sum_{k=0}^{j-1} \mu_{x+k}^{a\bullet} - j\delta\right) \frac{1 - \exp(-\delta - \mu_{x+j}^{a\bullet})}{\mu_{x+j}^{a\bullet} + \delta}. \end{aligned} \quad (3.12)$$

Second, the accelerated death benefit payable at diagnosis is covered by  $\alpha \bar{A}_{x;n] }^{a;a \rightarrow i}$  where the expression for  $\bar{A}_{x;n] }^{a;a \rightarrow i}$  can be found in (3.3). The remaining part of death benefits involves the following integral

$$\begin{aligned} \int_0^n {}_t p_x^{aa} \mu_{x+t}^{ai} \int_0^{n-t} {}_z p_{x+t;0}^{ii} \mu_{x+t+z;z}^{id} v(0, t+z) dz dt \\ = \int_0^n {}_t p_x^{aa} \mu_{x+t}^{ai} v(0, t) \bar{A}_{x+t;n-t] }^{i;i \rightarrow d} dt \end{aligned} \quad (3.13)$$

where

$$\bar{A}_{x+t;n-t] }^{i;i \rightarrow d} = \int_0^{n-t} {}_z p_{x+t;0}^{ii} \mu_{x+t+z;z}^{id} v(t, t+z) dz \quad (3.14)$$

is the present value of a unit benefit payable at death before time  $n$  for an individual being diagnosed with cancer at time  $t$ . Then, (3.13) can be rewritten as

$$\sum_{j=0}^{n-1} \exp\left(-\sum_{l=0}^{j-1} \mu_{x+l}^{a\bullet} - j\delta\right) \int_0^1 \exp\left(-t(\mu_{x+j}^{a\bullet} + \delta)\right) \mu_{x+j}^{ai} \bar{A}_{x+j+t;n-j-t] }^{i;i \rightarrow d} dt, \quad (3.15)$$

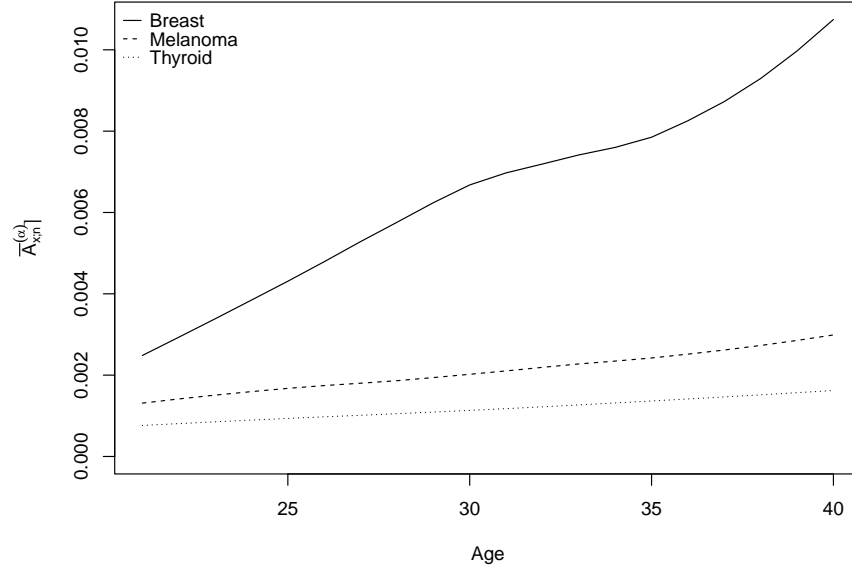
with the understanding that an empty sum is zero, where

$$\begin{aligned} \bar{A}_{x+j+t;n-j-t] }^{i;i \rightarrow d} &= \sum_{k=0}^{n-j-2} \exp\left(-\sum_{l=0}^{k-1} \tilde{\mu}(x+j+l;l)\right) \tilde{\mu}(x+j+k;k) v(j, j+k) \\ &\quad \int_0^1 \exp\left(-z(\tilde{\mu}(x+j+k;k) + \delta)\right) dz \\ &\quad + \exp\left(-\sum_{l=0}^{n-j-2} \tilde{\mu}(x+j+l;l)\right) \tilde{\mu}(x+n-1; n-j-1) v(j, n-1) \\ &\quad \int_0^{1-t} \exp\left(-z(\tilde{\mu}(x+n-1; n-j-1) + \delta)\right) dz \end{aligned} \quad (3.16)$$

and

$$\begin{aligned} \int_0^1 \exp\left(-z(\tilde{\mu}(x+j+k;k) + \delta)\right) dz \\ = \frac{1 - \exp\left(-\tilde{\mu}(x+j+k;k) - \delta\right)}{\tilde{\mu}(x+j+k;k) + \delta} \\ \int_0^{1-t} \exp\left(-z(\tilde{\mu}(x+n-1; n-j-1) + \delta)\right) dz \\ = \frac{1 - \exp\left(-(1-t)(\tilde{\mu}(x+n-1; n-j-1) + \delta)\right)}{\tilde{\mu}(x+n-1; n-j-1) + \delta}. \end{aligned} \quad (3.17)$$

Premiums  $\bar{A}_{x;n] }^{(\alpha)}$  are computed as a function of age at policy issue  $x \in \{20, 21, \dots, 40\}$  with  $n = 20$  and a yearly technical interest rate of 1%, that is,  $\delta = \ln 1.01$ . We take unit death benefit and  $\alpha = 50\%$ . The numerical values are displayed in Figure 3.7.



**Figure 3.7:** Values of  $\bar{A}_{x:n}^{(\alpha)}$  as function of age  $x$  for different cancer types with  $n = 20$ , yearly interest rate 1% and  $\alpha = 50\%$ .

We can see there that the product is cheaper for melanoma and thyroid cancers for all considered ages, thanks to lower incidence rates. Furthermore, higher premiums increasing with age are obtained for breast cancer, because of higher incidence rates.

**Remark 3.3.2.** A temporary life annuity starting at diagnosis can also be financed by “accelerating” the payment of (part of) the death benefit. Specifically, the temporary life annuity starting at diagnosis (that is, at entry in state  $i$ ) is payable continuously at rate  $b_i$ . Denoting as  $c_{ad}$  the amount of death benefit in case of transition from  $a$  to  $d$ , the residual amount of death benefit for an insured in state  $i$  dying after having spent a duration  $z$  in state  $i$  is given by

$$c_{id}(t, z) = \max\{c_{ad} - b_i z, 0\} = (c_{ad} - b_i z)_+. \quad (3.18)$$

Setting the duration payment  $m$  equal to  $c_{ad}/b_i$  and converting the death benefit into a temporary life annuity starting at diagnosis, the expected present value of insurance benefits is given by

$$\begin{aligned} & c_{ad} \bar{A}_{x:n}^{a;a \rightarrow d} + b_i \bar{a}_{x; c_{ad}/b_i}^{ai} \\ & + \int_0^n {}_t p_x^{aa} \mu_{x+t}^{ai} \left( \int_0^{c_{ad}/b_i} (c_{ad} - b_i z) {}_z p_{x+t;0}^{ii} \mu_{x+t+z;0}^{id} v(0, t+z) dz \right) dt. \end{aligned} \quad (3.19)$$

### 3.3.4 Cover option

Property loans are often accompanied with mortgage insurance that pays the balance of the loan if the mortgagor dies. Coverage is usually awarded in the form of term insurance with decreasing sum insured, with the amount of death benefit diminishing as the debt is reimbursed. This is common practice in France and Belgium. If the

insurer refuses to cover the risk of premature death then the bank does not lend the money, resulting in a barrier to property and home ownership (in case of house loan) and to entrepreneurship (in case of professional loan).

The product considered in this section is issued in state  $a$  and offers the beneficiary the option to obtain mortgage insurance at standard conditions even if he or she has been diagnosed with cancer (subject to a deferred period  $f$ ). The contract stipulates the characteristics of the loan (amount borrowed, amortization plan, maximal loan-to-value of the acquired building, etc.), or puts some limits. The sum insured is the difference between the actual single premium and the reference single premium computed from XK life table. This can also be seen as the expected cost on the Belgian market of a decision like the one taken by Cr dit Mutuel in France. We restrict our analysis to single premiums to reduce risk for the insurer.

Let  $\Pi_{x+t}^{XK}$  be the reference single premium for mortgage insurance securing the loan described in the policy conditions, at age  $x + t$ . The actual premium, given the extra mortality related to cancer for a patient aged  $x + t$  who has been diagnosed at time  $t - z$  is denoted as  $\Pi_{x+t;z}^i$ . The sum insured is the difference between these two premiums. To avoid under-pricing, we consider that policyholders will exercise their option at the worst time for the insurer. This leads to the worst-case expected present value

$$\text{EPV}_{\text{wc}}(x, n, f) = \max_s \int_0^n {}_t p_x^{aa} \mu_{x+t}^{ai} {}_f p_{x+t}^{ii} \left( \Pi_{x+t+f+s}^i - \Pi_{x+t+f+s}^{XK} \right) v(0, t + s + f) dt \quad (3.20)$$

where  $n$  is the coverage period and  $f$  is the deferred period stipulated by the contract.

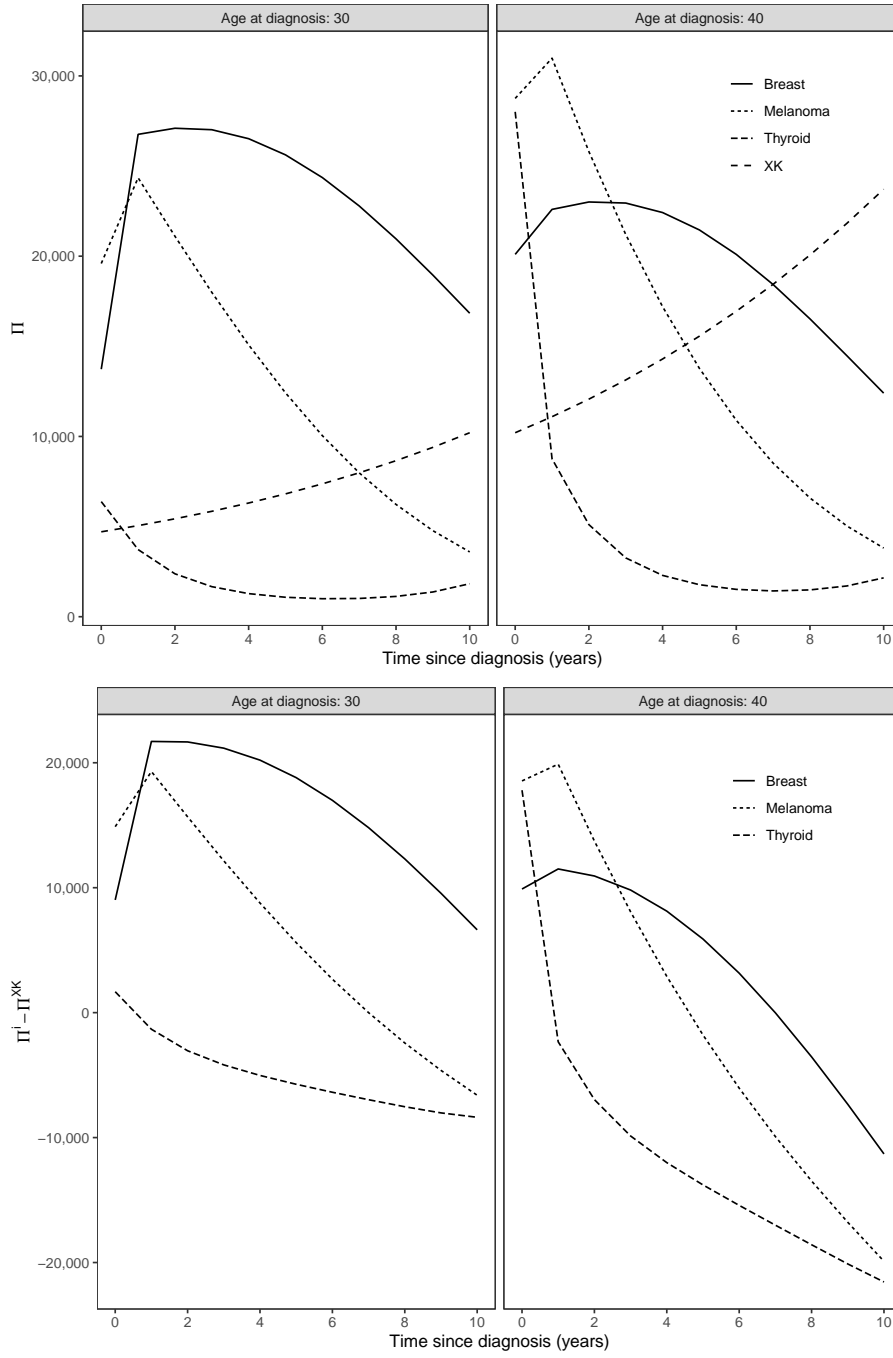
We consider here the same reference outstanding loan balance cover as in Soetewey et al. (2021). Specifically, we consider a home loan of duration 20 years (typical duration in Belgium). The mortgage insurance applicant aged  $x$  borrows an amount 100,000 at interest rate 2%. The technical interest rate for mortgage insurance is 1% and the insurance cover is over the full term of 20 years. These characteristics have been chosen as they represent a rather standard setting.

It is well documented that excess mortality generally decreases with time since diagnosis for most cancer types. Figure 6 in Soetewey et al. (2021) shows that  $\Pi_{x+t;z}^i$  peaks at  $z = 0$  for thyroid cancer at ages 30 and 50 and melanoma at age 50 or  $z = 1$  for melanoma at age 30 before decreasing for larger values of  $z$ . The upper panel in Figure 3.8 displays premiums  $\Pi_{x+t+k}^i$  and  $\Pi_{x+t+k}^{XK}$  for  $x + t = 30$  and 40, and time  $k$  since diagnosis in  $\{0, 1, \dots, 10\}$ . The corresponding differences are shown in the lower panel in Figure 3.8. When the difference is negative, the cover described in this section is not needed since cancer patients can be covered at standard premium rates. We can see there that the maximum is generally attained at  $s = 0$  when  $f \geq 2$ .

Let us take  $f = 2$  so that access to mortgage insurance at standard rate is granted two years after diagnosis. This is expected to address patients' needs since it is very unlikely that they wish to buy a house right after diagnosis, the two-year deferred period being devoted to the acute phase of treatment. Notice that this can be offered at no cost for thyroid cancer since  $\Pi_{x+t;t}^i$  falls below  $\Pi_{x+t}^{XK}$  two years after diagnosis in that case. The worst-case expected present value is then equal to

$$\begin{aligned} \text{EPV}_{\text{wc}}(x, n, 2) &= \sum_{j=0}^n v(0, j) {}_j p_x^{aa} \\ &\quad \int_0^1 {}_t p_{x+j}^{aa} \mu_{x+j}^{ai} {}_2 p_{x+j+t}^{ii} v(j, j + t + 2) \left( \Pi_{x+j+2}^i - \Pi_{x+j+2}^{XK} \right) dt \end{aligned} \quad (3.21)$$

where we have assumed that insurance premiums only depend on integer age. When



**Figure 3.8:** Values of premiums  $\Pi_{x+t+k;k}^i$  and  $\Pi_{x+t+k}^{XK}$  for  $x+t=30$  and  $40$ , and time  $k$  since diagnosis in  $\{0, 1, \dots, 10\}$ , with  $n=20$ , yearly interest rate 1% and the reference outstanding balance cover.

transition intensities are piecewise constant, we get

$$\begin{aligned}
& \sum_{j=0}^{n-1} \exp \left( - \sum_{l=0}^{j-1} \mu_{x+l}^{a\bullet} - j\delta \right) \mu_{x+j}^{ai} \exp(-2\delta) \\
& \quad \int_0^1 \exp \left( -t(\mu_{x+j}^{a\bullet} + \delta) \right) \exp \left( -\tilde{\mu}(x+j;0) - \tilde{\mu}(x+j+1;1) \right) \left( \Pi_{x+j+2;2}^i - \Pi_{x+j+2}^{KK} \right) dt \\
& = \sum_{j=0}^{n-1} \exp \left( - \sum_{l=0}^{j-1} \mu_{x+l}^{a\bullet} - j\delta \right) \mu_{x+j}^{ai} \exp(-2\delta) \\
& \quad \frac{1 - \exp \left( -\mu_{x+j}^{a\bullet} - \delta \right)}{\mu_{x+j}^{a\bullet} + \delta} \exp \left( -\tilde{\mu}(x+j;0) - \tilde{\mu}(x+j+1;1) \right) \left( \Pi_{x+j+2;2}^i - \Pi_{x+j+2}^{KK} \right).
\end{aligned} \tag{3.22}$$

The upper panel in Figure 3.9 displays the sum insured  $\Pi_{x+t+2;2}^i - \Pi_{x+t+2}^{KK}$  according to age  $x+t \in \{20, 21, \dots, 50\}$ . We can see that the differences are negative for thyroid cancer so that the product is not needed in that case, as patients can be covered at standard rates after the deferred period. Amounts  $\text{EPV}_{\text{wc}}(x, n, 2)$  are displayed in the lower panel of Figure 3.9 for  $x \in \{20, 21, \dots, 40\}$ , with  $n = 20$ , yearly interest rate 1% and the reference outstanding balance cover. Only melanoma and breast cancer are considered since the cover is not relevant for thyroid cancer. The cost appears to be moderate for melanoma but much higher for breast cancer. This results again from the higher incidence of breast cancer compared to melanoma.

### 3.4 Discussion

In this paper, we have developed a Semi-Markov 3-state model for designing and pricing cancer insurance products on a market where the “right to be forgotten” has been implemented. Different covers are proposed and three cancer types are considered for illustration, with different incidence rates and survival prognosis. It is shown that insurance products can be developed to address the particular needs of patients during the waiting period opening the “right to be forgotten”, but that costs greatly vary according to cancer type.

Insurance covers are typically limited to the first cancer occurrence. Several types of cancer can thus easily be combined into a single model to design products offering protection against more than just one cancer site. Considering the three cancer types considered in this paper, this would result in a hierarchical Semi-Markov model with 5 states, with state  $i$  replaced with three states  $i_1, i_2$  and  $i_3$  corresponding respectively to thyroid, melanoma and breast cancer, with transitions from state  $a$  to states  $\{i_1, i_2, i_3, d\}$  and from each  $i_1, i_2$  and  $i_3$  to state  $d$ . The extension to this more general setting is thus straightforward. Of course, distinguishing among cancer types is only relevant if coverage conditions vary with cancer site (for premium calculation, since distinguishing cancer types provides the actuary with a better understanding of cash flows, for instance to compute accurate reserves once a claim has been filed).

In Belgium, new diseases like, amongst others, HIV, some types of hepatitis and leukemia qualified for the “right to be forgotten”. Even if the present paper restricts to cancer insurance, a more general critical illness approach would theoretically also be possible. The main difficulty however is the lack of nationwide registry for these diseases. An appropriate source of reliable and representative data must thus be identified to perform actuarial calculations assessing the actual costs of these extensions to the “right to be forgotten”.



**Figure 3.9: Differences  $\Pi_{x+t+2,2}^i - \Pi_{x+t+2}^{XK}$  according to age  $x+t \in \{20, 21, \dots, 50\}$  in the upper panel and  $EPV_{wc}(x, n, 2)$  for  $x \in \{20, 21, \dots, 40\}$ , with  $n = 20$ , yearly interest rate 1% in the lower panel.**





# Health indices for disease incidence risk and duration in the Semi-Markov setting

## 4

This chapter corresponds to an article published jointly with Catherine Legrand, Michel Denuit and Geert Silversmit, carrying the same name as the chapter and published in the *xxx* in ?.

### **Abstract**

xxx add abstract

*Keywords:*



# Right to be forgotten for mortgage insurance issued to cancer survivors: Critical assessment and new proposal

## 5

This chapter corresponds to an article published jointly with Catherine Legrand, Michel Denuit and Geert Silversmit, carrying the same name as the chapter and published in the xxx in ?.

### Abstract

Soetewey et al. (2021) proposed to determine the waiting period opening the right to be forgotten (RTBF) as the time after diagnosis needed for the premium to revert back to some acceptable level expressed by means of regulatory life tables. However, this approach requires data up to 30 years after diagnosis (10 years of standard RTBF plus the typical duration of the loan), or extrapolating the results up to that time horizon. When survival statistics are only available over a shorter duration, it turns out that the results may strongly depend on the extrapolation method. This is why an alternative method is proposed here, based on a constraint imposed to the premium. This constraint is then transposed into a target on the conditional observed survival and the waiting period follows. For the sake of robustness, results obtained with the proposed approach are compared to results obtained with Kaplan-Meier estimate taken as a non-parametric reference. Furthermore, the paper investigates the impact of the stage of the tumor at diagnosis on waiting periods.

*Keywords:* Term insurance, impaired lives, waiting period, home loan, cancer stage at diagnosis.

### 5.1 Introduction and motivation

Outstanding balance insurance is generally required by lenders to secure their loans. The borrower is the insured life: if he or she dies before the loan has been fully repaid then the insurer pays a death benefit corresponding to the balance of the loan. As with any other term life insurance product, applicants with poor health conditions may be denied insurance or charged increased amounts of premium compared to standard conditions. In extreme cases, this may prevent them from accessing property or develop their business project. For this reason, several EU countries passed laws to ease access to mortgage insurance for long-term disease survivors.

The first initiative dates back to 2007, when France launched the AERAS Convention (AERAS is the acronym for “*s’Assurer et Emprunter avec un Risque Aggravé de Santé*” in French, which could be translated as “insuring and borrowing under poor health conditions”). This agreement, signed by the public authorities, banking and insurance sectors, and patients’ and consumers’ associations purposed to allow people cured from cancer or suffering certain chronic diseases to access insurance comprising benefits in case of death or disability, as well as to guaranteed income insurance. Considering long-term cancer survivors, the AERAS Convention included a “right to be forgotten” (henceforth abbreviated as RTBF), that is, the right for an insurance applicant not to declare a previous cancer after a period of 10 years starting at the end of the therapeutic protocol (reduced to 5 years for pediatric cancers). These periods of 10 and 5 years start from the date of the end of the therapeutic treatment, in absence of relapse within this period.

In Belgium, it was only in 2019 that the RTBF entered the insurance law. Based to a large extent on the reference tables published in the AERAS Convention, a Royal Decree dated May 26, 2019 lists certain types of cancer for which, depending upon entry criteria (such as cancer stage or age), the standard waiting period of 10 years from the end of active treatment is reduced. The RTBF has recently been adapted in Belgium, again following similar changes in France. As from November 2022, the standard waiting period opening the RTBF has been shortened from 10 years to 8 years, and it has been adopted that it will be reduced to 5 years as from January 2025. Moreover, also as from November 2022, the period is shortened to 5 years for cancer survivors who have been cured before the age of 21.

The RTBF has now also been installed in Luxembourg, The Netherlands, Portugal and Romania. It is being debated and advocated for at the European level to expand to the other EU countries as well. There are some ongoing discussions between Insurance Europe, the European Commission and the European Parliament on a possible EU-wide RTBF for cancer survivors. See for instance Scocca and Meunier (2020, 2022).

This paper aims to contribute to this evolution by proposing an actuarially sound methodology to evaluate a technically correct waiting period opening the RTBF. It starts with a critical assessment of the approaches proposed by Soetewey et al. (2021) and by Van Ginckel et al. (2022). It turns out that the results obtained from the method proposed by Soetewey et al. (2021) strongly depend on the extrapolation method. This is precisely shown in this paper, by modifying the extrapolation method and ending up with different waiting periods for some cancer types. This is clearly not acceptable in the context of the RTBF. To be precise, the problem is not with the method proposed by Soetewey et al. (2021) but comes from the limited follow-up period for patients in some cancer registries, including the Belgian one. This requires extrapolation to longer times since diagnosis and this step may induce higher uncertainty. If the length of the follow-up is enough, the method proposed by Soetewey et al. (2021) remains actuarially sound.

We also consider the approach proposed by Van Ginckel et al. (2022), which applies a biostatistical approach based on an arbitrary cut-off of 0.99 for the conditional net survival to propose reduced waiting periods for breast cancer. In Section 5.4.2, we show that, despite the apparent closeness to general population mortality in terms of survival, their method results in one-year death probabilities up to 10 times higher compared to general population at young adult ages. It is clear that such excess mortality cannot be absorbed by mortgage insurance market without increasing premiums at standard conditions.

There is thus a need for another approach, escaping the problem faced with extrapolation in case of limited follow-up and controlling the resulting premium compared to some market reference. In this paper, we impose a constraint to the premium

and transpose it into a target on the conditional observed survival probabilities. The main assumption retained throughout this paper is that mortality for cancer patients temporarily peaks after diagnosis before reverting back to a level comparable to the general population for survivors. We will demonstrate in this paper how the length of the waiting period opening the RTBF can be derived from the comparison of the conditional one-year survival probabilities of cancer patients with the corresponding probabilities at general population level. The main advantage is that the time from which the RTBF can be exercised can be estimated from the available data only, without the need to extrapolate mortality rates beyond 10 years. While cancer stage at diagnosis has not been taken into account in Soetewey et al. (2021), the present paper studies how the length of the waiting period opening the RTBF varies according to the extent of the tumor at diagnosis.

The remainder of this paper is structured as follows. Section 5.2 describes the mortgage insurance product considered in this paper. Section 5.3 presents the data used to perform the present study. Section 5.4 critically assesses the methods proposed by Soetewey et al. (2021) and by Van Ginckel et al. (2022). It is shown that the chosen extrapolation method for limited follow-up impacts on the length of the resulting waiting period opening the RTBF. Our alternative approach is detailed in Section 5.5, and results obtained with our approach are compared with results obtained via a method based on the Kaplan-Meier non-parametric estimator to demonstrate that the proposed approach is trustworthy. Waiting periods are then derived from the comparison between cancer patients' conditional one-year survival probabilities and conditional one-year survival probabilities of the general population. In Section 5.6, we illustrate these analyses considering Belgian data on melanoma, thyroid and female breast cancers according to the stage of tumor at diagnosis. The final Section 5.7 concludes the paper with a discussion.

## 5.2 Mortgage insurance

The proposed approach is based on a representative mortgage insurance contract for the market under consideration. In this paper, we work with a simplified example which could be an appropriate starting point. Precisely, we consider a mortgage insurance applicant aged  $x$  borrowing an amount of capital  $\kappa$  at annual interest rate  $r$  for a duration  $n$ . The capital is reimbursed by constant yearly installments over the  $n$  years. The borrower pays the amount

$$\frac{\kappa}{a_{\bar{n}|}}, \text{ where } a_{\bar{n}|} = \sum_{k=1}^n \frac{1}{(1+r)^k},$$

back to the lender. Working with annual repayments compared to monthly ones is conservative from the insurer's point of view.

At time  $s$ , the amount of the loan that has not yet been amortized is denoted as  $c_s$ . Let  $\lfloor s \rfloor$  denotes the integer part of  $s \in [0, n]$ , that is, the largest integer smaller than, or equal to  $s$ . At time  $s$ ,  $0 < s \leq n$ , the present value of future payments is

$$c_s = c_{\lfloor s \rfloor} (1+r)^{s-\lfloor s \rfloor}$$

where  $c_{\lfloor s \rfloor}$  is the outstanding balance of the loan at time  $\lfloor s \rfloor$ , right after the yearly installment has been paid, given by

$$c_{\lfloor s \rfloor} = \kappa \frac{a_{\overline{n-\lfloor s \rfloor}|}}{a_{\bar{n}|}}.$$

The loan is secured by a mortgage insurance, repaying the lender the amount  $c_s$  in case the policyholder dies at time  $s$ ,  $0 < s \leq n$ . The net single premium is the expected

present value (henceforth abbreviated as EPV) of insurance benefits, that is,

$$\pi_0 = \int_0^n {}_s p_x \mu_{x+s} c_s (1+i)^{-s} ds$$

where  $i$  is the technical interest rate for the insurance contract,  ${}_s p_x$  is the  $s$ -year survival probability for a policyholder aged  $x$  at policy issue and  $\mu_{x+s}$  is the hazard rate, or force of mortality, at attained age  $x+s$ . Note that hazard rate and force of mortality are the same. For consistency, from now on, hazard rate will always be used in this paper. In accordance with actuarial notation, we denote as  $p_y$  the one-year survival probability at integer age  $y$  (that is, the probability of being alive at age  $y+1$  given that the individual is alive at age  $y$ ). Note that  $y$  is introduced to differentiate from  $x$  which refers to the age at policy issue. In this sense, although both are equal in terms of value,  $y$  refers to a generic age while  $x$  refer to the specific age at policy issue.

Premium calculation is often based on regulatory or experience life tables. In this paper, we consider that standard conditions correspond to premiums computed according to the Belgian regulatory life table XK applying to insurance products comprising benefits in case of death (formally, XK defines minimum premium amount for policies with a positive sum at risk). This life table is widely adopted by Belgian insurers. It is known to be conservative and to generate a relatively high safety loading. Insurers are also allowed to use experience life tables available from the website of the National Bank of Belgium (NBB). These life tables reflect the mortality observed on the market, within portfolios of companies controlled by NBB. There is no safety loading and insurers are only allowed to apply premium rates resulting from NBB tables for relatively short periods of time (5 years, and then rates are subject to revision in case the observed mortality on the market changes over time). In this paper, we only consider the XK life table for premium calculation since these tables can be guaranteed for the whole contract duration and their conservatism better reflects increased mortality levels due to the disease.

The XK life table published in a Royal Decree does not distinguish between male and female policyholders, in accordance with EU anti-discrimination directive. For this reason, the entire analysis is conducted in this paper by pooling male and female mortality data. Also, it only gives one-year survival probabilities  $p_y$  at integer ages  $y$ . In this paper, we work under piecewise constant hazard rate, assuming that

$$\mu_{y+s} = \mu_y = -\ln p_y \text{ for all } 0 \leq s < 1 \text{ and integer } y.$$

Let us compute  $\pi_0$  under this assumption. To this end, we split the integral to get

$$\begin{aligned} \pi_0 &= \sum_{k=0}^{n-1} \int_k^{k+1} {}_s p_x \mu_{x+s} c_s (1+i)^{-s} ds \\ &= \sum_{k=0}^{n-1} {}_k p_x \int_0^1 {}_s p_{x+k} \mu_{x+k+s} c_k (1+r)^s (1+i)^{-k-s} ds \\ &= \sum_{k=0}^{n-1} {}_k p_x (1+i)^{-k} c_k \mu_{x+k} \int_0^1 {}_s p_{x+k} (1+r)^s (1+i)^{-s} ds. \end{aligned}$$

Now,

$$\begin{aligned} \int_0^1 {}_s p_{x+k} (1+r)^s (1+i)^{-s} ds &= \int_0^1 \exp\left(-s(\mu_{x+k} - \ln(1+r) + \ln(1+i))\right) ds \\ &= \frac{1 - \exp\left(-\mu_{x+k} \frac{1+r}{1+i}\right)}{\mu_{x+k} - \ln(1+r) + \ln(1+i)}, \end{aligned}$$

so that we finally get

$$\pi_0 = \sum_{k=0}^{n-1} {}_k p_x (1+i)^{-k} c_k \mu_{x+k} \frac{1 - \exp(-\mu_{x+k})^{\frac{1+r}{1+i}}}{\mu_{x+k} - \ln(1+r) + \ln(1+i)} \quad (5.1)$$

where  ${}_0 p_x = 1$  and for  $k \geq 1$ ,

$${}_k p_x = \prod_{j=0}^{k-1} p_{x+j} = \exp\left(-\sum_{j=0}^{k-1} \mu_{x+j}\right).$$

### 5.3 Data

The data available from the Belgian Cancer Registry (BCR) are considered in this paper. The BCR is a national population-based cancer registry collecting data on all new cancer cases diagnosed in Belgium since the incidence year 2004. Cancer registration has been made compulsory by law since 2006 in Belgium. The vital status is derived from linkage with the Belgian Crossroads Bank for Social Security up to April 11, 2022 and quality controls are performed regularly by BCR, ensuring the continuity and completeness of cancer registration in the country. More information can be found on the BCR website, at [www.kankerregister.org](http://www.kankerregister.org).

To illustrate our work, three cancer types are considered: melanoma (ICD-10 C43), thyroid (ICD-10 C73) and female breast (ICD-10 C50) cancer (only female breast cancer is considered as there are too few registrations for male breast cancer). These three cancer sites have been selected to evaluate the proposed method in different scenarios. Melanoma and thyroid cancer patients are known to have a limited excess hazard compared to the general population (Soetewey et al., 2021). The situation for female breast cancer patients is different with usually a high yearly survival probability in the first years after the date of diagnosis before it eventually decreases due to late cancer recurrences. Moreover, it is known that mortality for patients diagnosed with any of these three cancer types varies with time since diagnosis (Soetewey et al., 2022), yielding appropriate illustrations of the right to be forgotten.

For these applications, our analyses are also limited to patients aged 20 to 69 at time of diagnosis for two main reasons. First, childhood cancers can be seen as a category of cancer on their own, and are often studied separately because they greatly differ from adult cancers. Second, the RTBF mainly concerns young adults and active life.

Out of a total of 161,007 tumors, melanoma, thyroid and breast cancer represent, respectively, 29,213 (18.1%), 12,241 (7.6%) and 119,553 (74.3%) cases diagnosed between 2004 and 2020. Patients were followed-up until April 11, 2022, resulting in a follow-up ranging from 2 to 18 years. Only one record per patient (with the earliest incidence date) within each cancer site was kept for patients with multiple primary diagnoses. A minority of patients without national security number were excluded from the analysis. Patients lost to follow-up (mostly due to moving abroad) and patients still alive at the end of the follow-up period were treated as censored observations. Censoring is assumed to be uninformative.

Table 5.1 summarizes the number of included cases, number and proportion of deaths and percentage of lost to follow-up before April 11, 2022 per type of cancer, sex and age group. The fraction of patients lost to follow-up per subgroup varied from 1.31% for women with breast cancer aged 50-69 to 4.1% for male thyroid cancer patients aged 20-34. The total fraction of patients lost to follow-up cases, regardless of sex, site or age group was 1.64%. Moreover, mean age at diagnosis was 50.5 years (standard deviation 12.1), 48.1 years (standard deviation 12.4) and 54.6 years (standard deviation 9.5) for melanoma, thyroid and breast cancer, respectively.

Sex	Cancer site	Age at diagnosis	Lost to follow-up	Number of included cases	Number of deaths
Men	Melanoma	20-34	3.72%	969	94
		35-49	2.66%	3,266	404
		50-69	1.70%	7,460	1,583
	Total			11,695	2,081
Men	Thyroid	20-34	4.10%	366	6
		35-49	3.12%	961	67
		50-69	2.14%	1,773	379
	Total			3,100	452
Women	Melanoma	20-34	3.62%	2,488	78
		35-49	1.47%	6,137	382
		50-69	1.35%	8,893	1,112
	Total			17,518	1,572
Women	Thyroid	20-34	3.80%	1,607	14
		35-49	2.67%	3,449	107
		50-69	2.06%	4,085	484
	Total			9,141	605
Women	Breast	20-34	2.76%	3,112	502
		35-49	1.78%	32,743	4,058
		50-69	1.31%	83,698	15,946
	Total			119,553	20,506

**Table 5.1: Number of persons diagnosed with melanoma, thyroid and female breast cancer in Belgium between 2004 and 2020 (BCR data) by sex, site and age group, together with the percentage of lost to follow-up and the number of deaths.**



For the mortality in the general population, Belgian population life tables are available from Statbel (the Belgian statistical office) and can be freely downloaded from the website [www.statbel.fgov.be](http://www.statbel.fgov.be).

## 5.4 Critical assessment

In this section, we revisit previous studies by Soetewey et al. (2021) and Van Ginckel et al. (2022) to underline their possible shortcomings.

### 5.4.1 Impact of extrapolation in case of limited follow-up

In this paper, we analyze survival time from diagnosis for cancer patients according to a number of covariates summarized into the vector  $\mathbf{z}$ . Specifically,  $T$  denotes the remaining lifetime at diagnosis, so time from diagnosis to death. Given  $\mathbf{z}$ ,  $T$  has probability density function  $f(\cdot|\mathbf{z})$ , distribution function  $F(\cdot|\mathbf{z})$ , survival function  $S(\cdot|\mathbf{z}) = 1 - F(\cdot|\mathbf{z})$ , and hazard rate  $\lambda(\cdot|\mathbf{z}) = f(\cdot|\mathbf{z})/S(\cdot|\mathbf{z})$ . Contrarily to insurance studies,  $T$  denotes the remaining lifetime after diagnosis and age at diagnosis is included as a covariate (attained age is thus obtained by summing age at diagnosis and survival time). The link with the international actuarial notation for survival probabilities and hazard rate is as follows: if the insurance applicant aged  $x$  has been diagnosed with cancer at age  $x - w$  then

$$\begin{aligned} {}_s p_x &= \frac{S(w + s | \text{age at diagnosis} = x - w)}{S(w | \text{age at diagnosis} = x - w)} \\ \mu_{x+s} &= \lambda(w + s | \text{age at diagnosis} = x - w). \end{aligned}$$

Relative survival provides a measure of the excess mortality experienced by cancer patients by comparing the mortality in the cancer population with the mortality in the general population. Relative survival models are divided into additive and multiplicative models. Despite the wide acceptance of multiplicative specifications within the actuarial community, additive models are generally applied in cancer studies. The additive specification is thus adopted in this paper. The hazard rate  $\lambda(t|\mathbf{z})$  at time  $t$  since diagnosis for cancer patients with covariate vector  $\mathbf{z}$  is decomposed into two additive components: the population hazard based on available patient's characteristics  $\mathbf{z}$ , denoted as  $\lambda_P(t|\mathbf{z})$ , and the excess hazard specific for the disease of interest, denoted as  $\lambda_E(t|\mathbf{z})$ . Formally,

$$\lambda(t|\mathbf{z}) = \lambda_P(t|\mathbf{z}) + \lambda_E(t|\mathbf{z}). \quad (5.2)$$

In (5.2),  $\lambda_P(\cdot|\mathbf{z})$  usually corresponds to general population life tables. Here, the covariate vector  $\mathbf{z}$  corresponds to age and it is the same in  $\lambda_P(t|\mathbf{z})$  and  $\lambda_E(t|\mathbf{z})$ .

Soetewey et al. (2021) adopted the flexible parametric model proposed by Remontet et al. (2019) and Fauvernier et al. (2019a,b) to (i) allow for a flexible modeling of the baseline excess hazard, (ii) account for non-linear and non-proportional effects of covariates and (iii) allow for a flexible interaction between several covariates adopting a multidimensional penalized splines approach. This leads to the specification

$$\ln \lambda_E(t|\mathbf{z}) = \sum_{j=1}^J g_j(t, \mathbf{z}) \quad (5.3)$$

where  $g_j(\cdot, \cdot)$  are uni- or multidimensional penalized spline function. This model has the advantage that the splines bring the flexibility needed for modeling the hazard and the penalty terms control this flexibility for smooth estimation. Excess hazard

Cancer site	Age at diagnosis	Waiting period (in years)	
		Soetewey et al. (2021)	Alternative extrapolation method
Melanoma	30	9	>10
Melanoma	50	3	4
Thyroid	30	1	1*
Thyroid	50	1	1
Breast	30	>10	>10
Breast	50	3	6

**Table 5.2: Waiting periods by cancer site and age at diagnosis. The star indicates that EPV does not stay below XK level but start to increase a few years after diagnosis.**

was estimated using the `flexrsurv` package in R (Clerc-Urmès and Grzebyk, 2023), assuming non-linear and non-proportional hazard for age at diagnosis.

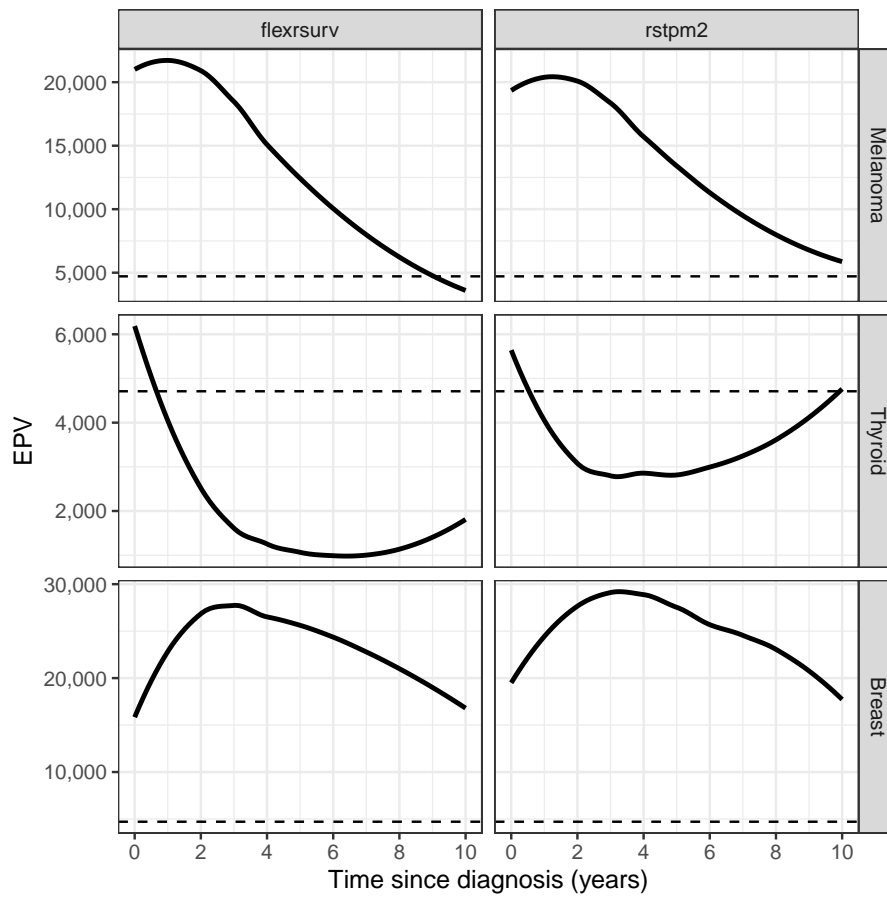
Let us proceed as in Soetewey et al. (2021) and determine the waiting period opening the RTBF as the smallest duration after diagnosis so that the expected present value of mortgage insurance benefits gets back to the premium determined according to XK life table. This is represented in Figures 5.1 and 5.2 for a patient diagnosed at the age of 30 and 50, respectively. The left panels are based on the estimation and extrapolation method adopted in Soetewey et al. (2021), implemented in the R package `flexrsurv`. The right panels are based on an alternative method implemented in the R package `rstopm2` (Jakobsen et al., 2020; Liu et al., 2017, 2018; Zhan et al., 2018). For this alternative method, we used a flexible parametric survival model with proportional hazards and 3 degrees of freedom for modelling the baseline log-cumulative hazard. These characteristics have been chosen to obtain excess hazards that are as similar as possible to the ones obtained with the `flexrsurv` package, and other scenarios revealed drastically different patterns for the two considered ages at diagnosis. The resulting waiting periods are listed in Table 5.2. They are obtained by considering that patients become insurable at standard conditions when the EPV reaches the level set by the XK life table. We can see in Table 5.2 that for melanoma and female breast cancer diagnosed at age 50, the waiting periods determined as in Soetewey et al. (2021) are smaller compared to the alternative extrapolation method. For thyroid cancer patients aged 30, the waiting period remains 1 year but EPV exceeds XK level a few years later according to the alternative extrapolation method. For melanoma cancer diagnosed at age 30, the reduced waiting period determined as in Soetewey et al. (2021) is contradicted by the alternative extrapolation method.

This example shows that conclusions may rely to a large extent on the extrapolation method, even more so when one considers different model parameters. This is not acceptable in the context of the RTBF. The aim of this paper is to propose a new approach, only using the available data (so without the need to extrapolate mortality rates beyond 10 years).

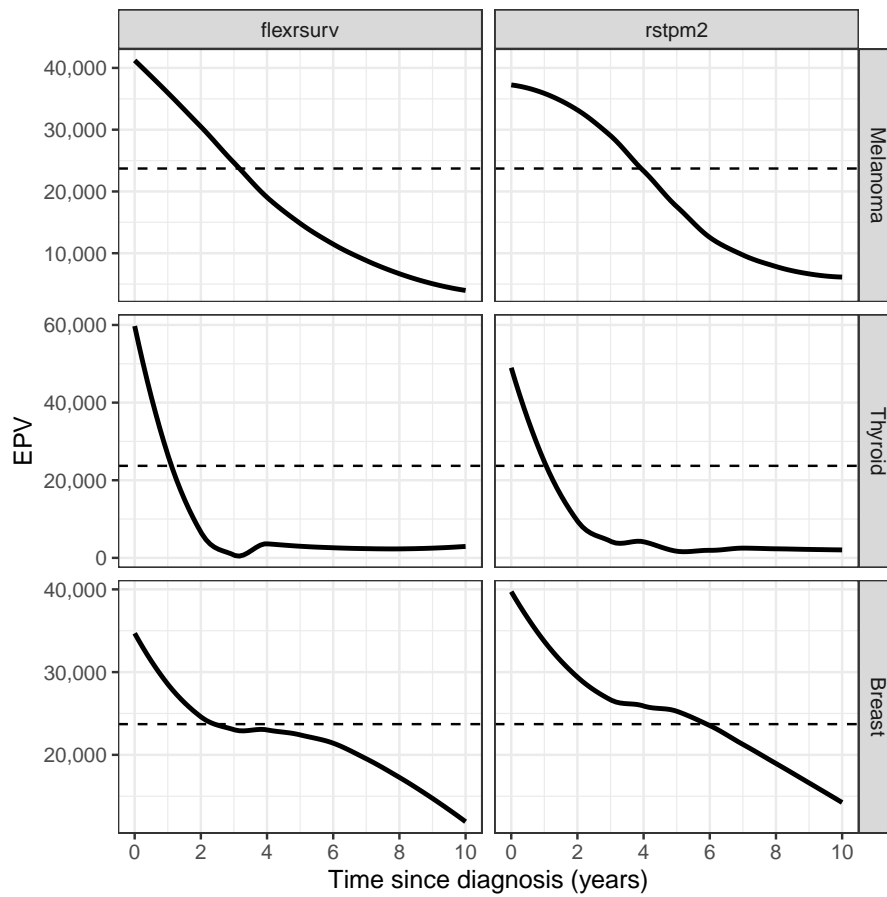
#### 5.4.2 Conditional relative net survival

Van Ginckel et al. (2022) applied a pure biostatistical approach based on an arbitrary cut-off of 0.99 for the conditional net survival to propose reduced waiting periods for breast cancer. This section explains why their apparently sound methodology fails to convince actuaries.

In accordance with actuarial notation, let  $q_y = 1 - p_y$  be the one-year death probability at age  $y$  (that is, the probability of dying before age  $y+1$  given that the individual is alive at age  $y$ ). Probabilities, corresponding to general population mortality, are henceforth denoted as  $p_y^{\text{NIS}}$  and  $q_y^{\text{NIS}}$  where “NIS” refers to the National Institute of



**Figure 5.1: Expected present value (EPV) of a life insurance contracted by a 30-year-old cancer patient for a period of 20 years with interest of 1 percent and benefit of 100 000. Horizontal dashed lines correspond to EPV calculated according to XK life table.**



**Figure 5.2: Expected present value (EPV) of a life insurance contracted by a 50-year-old cancer patient for a period of 20 years with interest of 1 percent and benefit of 100 000. Horizontal dashed lines correspond to EPV calculated according to XK life table.**

Statistics (Statbel based in Brussels; [www.statbel.fgov.be](http://www.statbel.fgov.be)). Likewise, denote as  $p_{x,w}^{\text{CR}}$  and  $q_{x,w}^{\text{CR}}$  these probabilities for an individual of age  $x$  who was diagnosed with cancer  $w$  years ago, so at age  $x - w$ . Here, “CR” refers to Cancer Registry established at national level.

The “conditional relative net survival” referred to in Section 4.2.1 of Van Ginckel et al. (2022) can be interpreted as the ratio  $p_{x,w}^{\text{CR}}/p_y^{\text{NIS}}$ . The reduced waiting period is then determined as the smallest  $w$  such that  $p_{x,w}^{\text{CR}}/p_y^{\text{NIS}} > 0.99$ . Their argument is that the resulting  $w$  ensures that surviving patients’ mortality is very close to the general population one. However, when translated into premium calculation, this rule turns out to produce large increases. Indeed, considering that patients can be covered at standard conditions once they have survived  $w$  years after diagnosis, with one-year survival probability

$$p_{x,w}^{\text{CR}} = 0.99p_y^{\text{NIS}} \quad (5.4)$$

means that

$$q_{x,w}^{\text{CR}} = 1 - p_{x,w}^{\text{CR}} = 1 - 0.99p_y^{\text{NIS}} = q_y^{\text{NIS}} + 0.01p_y^{\text{NIS}}.$$

Hence, the one-year death probability (driving the amount of premium for a one-year term insurance) is increased by 1% times the corresponding one-year survival probability. The impact of this rule greatly varies according to age  $x$ :

- if  $q_y^{\text{NIS}} = 0.001$  then this results in an actual one-year death probability

$$0.001 + 0.01 \times 0.999 = 0.01099$$

which means that the one-year death probability (and hence the yearly term insurance premium) is multiplied by 10, approximately.

- if  $q_y^{\text{NIS}} = 0.01$  then this results in an actual one-year death probability

$$0.01 + 0.01 \times 0.99 = 0.0199$$

which means that the one-year death probability (and hence the yearly term insurance premium) is multiplied by 2, approximately.

Considering the typical age range where mortgage insurance is sold, the rule retained by Van Ginckel et al. (2022) allows for mortality levels which largely exceed those corresponding to general population.

## 5.5 Proposed approach for limited follow-up

Clearly, the rule defining reduced waiting periods for the RTBF must be expressed in terms of premiums. The question about the RTBF centers on evaluating extra claim costs and sharing them among stakeholders in a fair and transparent way. This can only be achieved by computing actual premiums at age  $x$  in function of the time  $w$  elapsed since diagnosis, and by comparing them to the reference levels XK corresponding to regulatory expected costs.

Let  $\pi_0^{\text{XK}}$  be the amount of premium obtained from (5.1) when survival probabilities and death rates correspond to the XK life table. Then, formula (5.1) is used again to obtain the additive increase in mortality compared to the general population level. Precisely, the additive mortality shift  $\gamma$  is the unique positive root of the equation

$$\pi_0^{\text{XK}} = \sum_{k=0}^{n-1} \exp \left( - \sum_{j=0}^{k-1} \left( \mu_{y+j}^{\text{NIS}} + \gamma \right) \right) (1+i)^{-k} c_k \left( \mu_{y+k}^{\text{NIS}} + \gamma \right) \frac{1 - \exp \left( - \left( \mu_{y+k}^{\text{NIS}} + \gamma \right) \frac{1+r}{1+i} \right)}{\mu_{y+k}^{\text{NIS}} + \gamma - \ln(1+r) + \ln(1+i)}. \quad (5.5)$$

The solution is unique because the right-hand side of this equation is increasing in  $\gamma$  and the left-hand side is larger than the right-hand side when  $\gamma = 0$  because the XK life table is conservative. The solution is therefore such that  $\gamma > 0$ .

Following the idea of (5.4), we propose to define the waiting period opening the RTBF as the smallest  $w$  such that

$$p_{x,w}^{\text{CR}} = \exp(-\gamma) p_y^{\text{NIS}}. \quad (5.6)$$

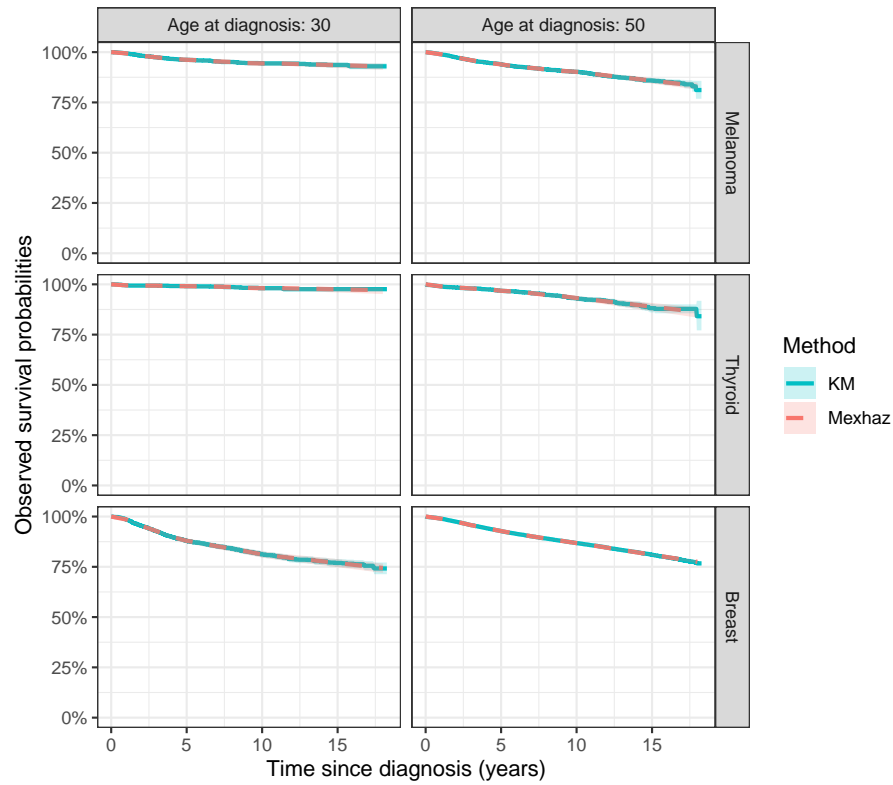
In this case, we recover a constraint on the conditional observed survival, but with the arbitrary 0.99 level replaced with  $\exp(-\gamma)$  controlling premium. Following (5.6), the waiting period opening the RTBF is determined as the smallest  $w$  such that  $p_{x,w}^{\text{CR}}/p_y^{\text{NIS}} > \exp(-\gamma)$ . To apply this rule,  $p_y^{\text{NIS}}$  can easily be found within Belgian population life tables, available from Statbel. The calculation of  $p_{x,w}^{\text{CR}}$  is explained in Appendix A.

Let us now apply this method to get the length of the waiting period opening the RTBF. To this end, survival probabilities of cancer patients, obtained via a flexible parametric model (using the `mexhaz` R package (Charvat and Belot, 2021) and based on a baseline hazard specified as the exponential of B-splines of degree 2 with a knot at 2.5 years of follow-up), are first compared with the observed survival probabilities obtained with the non-parametric Kaplan-Meier (1958) estimator. The results are displayed in Figure 5.3. It can be seen from Figure 5.3 that observed survival curves obtained via a flexible parametric model and via the Kaplan-Meier estimator are very similar for all cases under consideration (i.e., for both ages at diagnosis and for all three cancers of interest).

Secondly, conditional one-year observed survival probabilities, obtained via a flexible parametric model (denoted  $p_{x,w}^{\text{CR}}$  and detailed in Section A) are compared with the conditional one-year observed survival probabilities obtained based on the Kaplan-Meier estimator (henceforth denoted as  $p_{x,w}^{\text{KM}}$ ) in Figure 5.4. Here, probabilities  $p_{x,w}^{\text{KM}}$  are computed with increments of 0.1 year and are referred as the KM-based method in the remainder of the text since these probabilities are computed based on the Kaplan-Meier estimator. The difference with a standard Kaplan-Meier estimator is that  $p_{x,w}^{\text{KM}}$  correspond to conditional one-year observed survival probabilities (instead of simply observed survival probabilities). In practice,  $p_{x,w}^{\text{KM}}$  are found by computing one-year survival probabilities using the `survival` R package (Terry M. Therneau and Patricia M. Grambsch, 2000), repeatedly for each subgroup of patients who survived at least 0, 0.1, 0.2, ..., 10 years since diagnosis. The advantage of computing  $p_{x,w}^{\text{KM}}$  this way is that the provided confidence intervals are usable, which is not the case if  $p_{x,w}^{\text{KM}}$  are computed by dividing the survival probability at a given time by the survival probability one year earlier). The goal of comparing  $p_{x,w}^{\text{CR}}$  with a counterpart based on a non-parametric reference such as the Kaplan-Meier estimator is to demonstrate that results obtained with the proposed approach are trustworthy.

Figure 5.4 shows that conditional one-year survival probabilities obtained via the flexible parametric model follow globally the same trend than the ones obtained via the KM-based method for all scenarios, except for the first year after diagnosis for patients diagnosed with melanoma cancer at age 50. We consider that conditional one-year observed survival probabilities are reasonably well estimated with our approach when compared to a non-parametric reference. Furthermore, for a given sample size, confidence intervals are narrower with our approach compared with the non-parametric Kaplan-Meier, a reason to prefer the new approach over the non-parametric reference.

To determine the waiting period opening the RTBF, conditional one-year survival probabilities obtained via the two approaches, that is,  $p_{x,w}^{\text{CR}}$  and  $p_{x,w}^{\text{KM}}$ , are divided by the conditional one-year survival probabilities in the general population, that is,  $p_y^{\text{NIS}}$ .



**Figure 5.3: Survival probabilities (with 95% confidence interval) by cancer site and age at diagnosis. Mexhaz method corresponds to the probabilities obtained via a flexible parametric model (dashed line), whereas KM-based method corresponds to the ones obtained based on the non-parametric Kaplan-Meier estimator (solid line).**

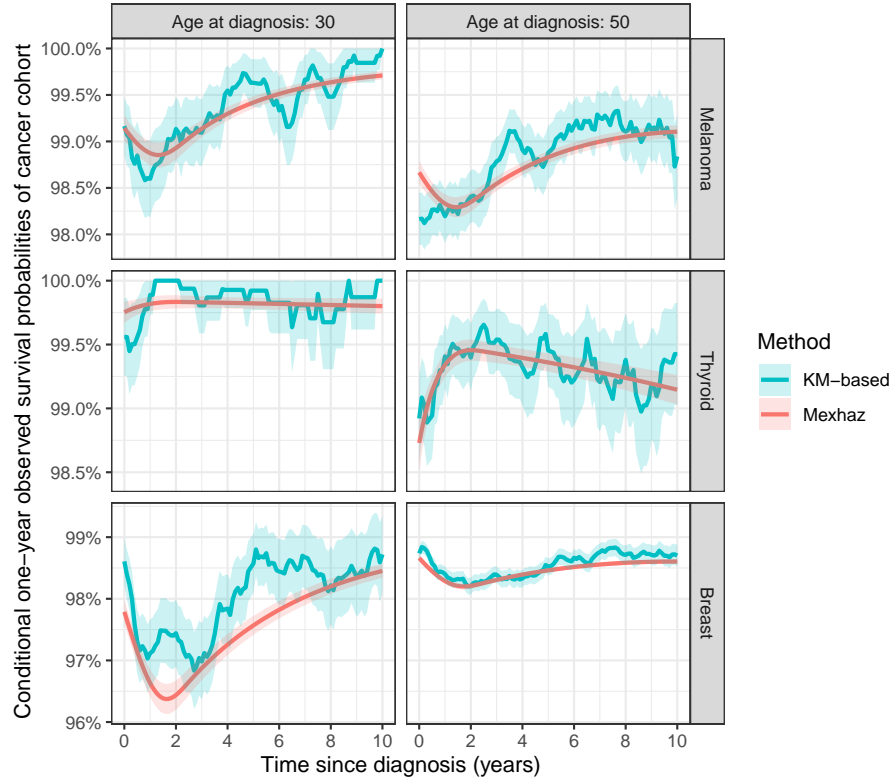


Figure 5.4: Conditional one-year survival probabilities (with 95% confidence interval) by cancer site and age at diagnosis. Mexhaz method corresponds to the probabilities obtained via a flexible parametric model,  $p_{x,w}^{CR}$ , whereas KM-based method corresponds to the ones obtained based on the non-parametric Kaplan-Meier estimator,  $p_{x,w}^{KM}$ .



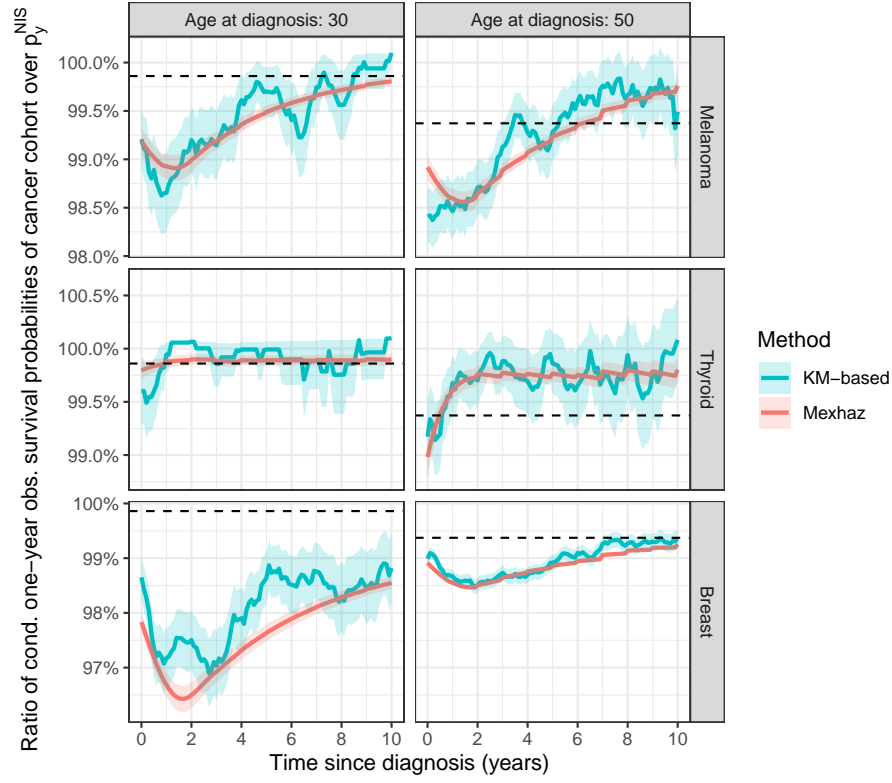
Cancer site	Age at diagnosis	Waiting period (in years)	
		Our approach	KM-based
Melanoma	30	> 10	10
Melanoma	50	6	6
Thyroid	30	1	Uncertain
Thyroid	50	1	2
Breast	30	> 10	> 10
Breast	50	> 10	> 10

**Table 5.3: Waiting periods by cancer site and age at diagnosis computed via our approach and via the KM-based method.**

Results are displayed in Figure 5.5. For the sake of comparison, the waiting period opening the RTBF is determined as the smallest  $w$  such that  $p_{x,w}^{CR}/p_y^{NIS} > \exp(-\gamma)$  or such that  $p_{x,w}^{KM}/p_y^{NIS} > \exp(-\gamma)$ . Dividing  $p_{x,w}^{CR}$  and  $p_{x,w}^{KM}$  by  $p_y^{NIS}$  allows the comparison with the additive correction  $\exp(-\gamma)$ . Here,  $\gamma = 0.0014$  at age 30 and  $\gamma = 0.0063$  at age 50. Notice the difference with the threshold of 0.99 set in Van Ginckel et al. (2022), as  $\exp(-\gamma)$  equals 0.998601 and 0.9937198 for a patient diagnosed at age 30 and 50, respectively. Also note that, for patients aged 30 years at diagnosis,  $p_{x,w}^{KM}$  is actually computed based on patients aged between 25 and 35 years at diagnosis. For patients aged 50 years at diagnosis,  $p_{x,w}^{KM}$  is computed based on patients aged between 45 and 55 years at diagnosis. This is to include more patients and thus have more stable estimates. Indeed, samples of patients diagnosed at exactly 30 and 50 years old have a limited size, in particular for thyroid cancer. Considering patients aged from 25 to 35 and from 45 to 55 instead of patients of exactly 30 and 50 years old does not undermine our analyses, as patients within each age group are very similar in terms of survival.

Results are displayed in Table 5.3 and Figure 5.5. Remember that an increasing ratio is a sign of better prognosis for cancer patients. On the other hand, a decreasing ratio is a sign that survival for cancer patients declines over the years since diagnosis, so a sign of worse prognosis compared to the general population. Following this, and in order to be as conservative as possible, if the ratio of conditional one-year survival probability reaches the level of the additive correction more than once within the 10-year period after diagnosis, the waiting period is set as the largest time after diagnosis where the ratio of survival probabilities crosses the additive correction level. Notice also the emergence of small jumps when plotting the ratio of the conditional survival probabilities in Figure 5.5 resulting from the division by the one-year survival probabilities  $p_y^{NIS}$  in the general population.

From Table 5.3 and Figure 5.5, we can see that waiting periods are below 10 years for melanoma cancer patients aged 50 years at diagnosis, and thyroid cancer patients aged 30 and 50 at diagnosis. Waiting periods obtained via the KM-based method are below 10 years for melanoma and thyroid cancer patients aged 50 at the time of diagnosis. For all other scenarios, waiting periods are equal or above 10 years after diagnosis. When comparing the two approaches, waiting periods are relatively equivalent for all considered subgroups except for thyroid cancer patients diagnosed at 30 years old, who have an uncertain waiting period via the KM-based method (since the ratios of conditional one-year survival probabilities fluctuate around the level set by the additive correction from 0 to 10 years after diagnosis). Moreover, breast cancer patients diagnosed at age 30 have a waiting period above 10 years while it is equal to 10 years according to the KM-based calculation. Recall that, as advocated in Soetewey et al. (2021), these waiting periods start at the time of diagnosis, and not at the end of



**Figure 5.5: Ratio of conditional one-year survival probability (with 95% confidence interval) by cancer site and age at diagnosis, together with the additive correction  $\exp(-\gamma)$  (horizontal dashed lines). Mexhaz method corresponds to  $p_{x,w}^{\text{CR}}/p_y^{\text{NIS}}$ , whereas KM-based method corresponds to  $p_{x,w}^{\text{KM}}/p_y^{\text{NIS}}$ .**

Cancer site	Tumor stage	Number of cases	Number of deaths	1-year survival prob. (95% CI)	5-year survival prob. (95% CI)
Melanoma	I	20,949	1,153	0.997 (0.996-0.997)	0.970 (0.968-0.973)
	II	3,019	777	0.980 (0.975-0.985)	0.802 (0.786-0.817)
	III	1,669	537	0.949 (0.939-0.960)	0.711 (0.688-0.735)
	IV	444	327	0.658 (0.615-0.703)	0.298 (0.257-0.345)
Thyroid	I	8,071	311	0.995 (0.994-0.997)	0.979 (0.976-0.982)
	II	946	73	0.989 (0.983-0.996)	0.961 (0.948-0.974)
	III	841	134	0.993 (0.987-0.999)	0.942 (0.926-0.958)
	IV	617	294	0.779 (0.747-0.812)	0.625 (0.587-0.665)
Breast	I	56,039	4,813	0.996 (0.995-0.996)	0.965 (0.964-0.967)
	II	37,935	5,979	0.992 (0.991-0.993)	0.926 (0.923-0.929)
	III	11,919	3,922	0.976 (0.973-0.979)	0.805 (0.798-0.813)
	IV	5,639	3,839	0.829 (0.819-0.839)	0.390 (0.377-0.404)

**Table 5.4: Number of included cases, number of observed deaths, one-year and 5-year observed survival probabilities (with 95% confidence interval) by cancer site and stage of the tumor. Survival probabilities are obtained with the non-parametric Kaplan-Meier estimator.**

the therapeutic protocol as it is the case with the current legislation.

## 5.6 Impact of the stage of the tumor

One could argue that mortality and thus the waiting period opening the RTBF varies between cancer patients diagnosed at different tumor stages. As this information is available in BCR, this section refines the preceding analyses by cancer stage at diagnosis.

Information on both the clinical and pathological staging has been combined to define a final tumor stage. First, clinical staging is an estimate of the extent of the cancer based on results of physical exams, imaging tests, endoscopy exams, biopsies, and for some cancers, the results of other tests, such as blood tests. Second, the pathological staging (also called the surgical stage) is an estimate of the extent of the cancer that is based on the results of pathological examination of the resection piece after surgery. In some cases, the pathological stage is different from the clinical stage, for instance, if the surgery shows the cancer has spread more than was seen on imaging tests. A common practice is to combine these two methods to obtain a so-called combined stage. When the pathological stage is known, it is taken as combined stage, unless there is clinical evidence of metastasis. In case the pathological stage is unknown, the clinical stage is retained. Combining the clinical and pathological stage limits missing values (missing combined stage appears only when both the clinical and pathological stages are missing). This combined stage is considered in this section.

Stages I, II, III and IV were considered. Tumors with an unknown stage at the time of diagnosis, representing 7.5% of all tumors, have been ignored. Number of included cases, number of observed deaths, one-year and 5-year observed survival probabilities (obtained with the non-parametric Kaplan-Meier estimator) by cancer site and stage of the tumor are displayed in Table 5.4. Given the small number of observations for stages III and IV, these two stages have been combined for the analyses. Furthermore, to ensure simplicity and given that cancer patients diagnosed at stages I and II are relatively similar in terms of survival, these two stages have also been combined.

The present section is aimed at studying the appropriateness of stratifying the

RTBF according to the stage of the tumor at diagnosis: waiting periods are computed separately for patients diagnosed at stages I–II and at stages III–IV using our proposed approach. This will serve as a comparison with results obtained before, where all stages are included. Note that, as the additive correction  $\exp(-\gamma)$  depends only on age at diagnosis, it differs between patients diagnosed at 30 and 50 years old, but it is the same for all stages and it remains the same than when including all stages. Notice that the non-parametric Kaplan-Meier reference is no longer used because stratifying by stage reduces drastically the number of observations, in particular for stages III and IV. This rises the issue of the accuracy of the Kaplan-Meier estimator, and therefore reduces its usefulness in the context of the RTBF.

Results of the stratification by stage at diagnosis are displayed in Table 5.5 and Figure 5.6. Waiting period is lower for patients diagnosed at stages I–II compared to patients diagnosed at stages III–IV for all scenarios. In particular, compared to patients diagnosed at all stages, when including only patients diagnosed at stages I–II, waiting periods are reduced from 6 to 4 years for melanoma cancer patients aged 50, reduced from 1 to 0 year for thyroid cancer patients aged 50, and reduced from more than 10 years to 7 years for female breast cancer patients aged 50. For melanoma cancer patients aged 30, thyroid cancer patients aged 30 and breast cancer patients aged 30, waiting periods remain the same whether it is calculated by stage or for all stages combined. This shows that, for the three cancer sites considered, stratifying the analyses according to the stage has no impact on the waiting periods for patients diagnosed at the age of 30, but has an impact for patients diagnosed at the age of 50. This can be partly explained by the fact that, among patients diagnosed at a young age, a small proportion is diagnosed at stages III–IV. For instance, only 8.36% of patients aged 30 or below at the time of diagnosis are diagnosed at stages III–IV. Furthermore, we observe that the waiting period is above 10 years for patients diagnosed at stages III–IV for all scenarios.

Notice that 95% confidence intervals for stages I–II (Figure 5.6) are narrower than when all stages are considered (Figure 5.5), although the sample size is smaller when including only patients diagnosed at stages I–II. This is explained by the fact that, for the same sample size, standard errors for probabilities closer to 0% or 100% are smaller. Therefore, even though the sample size is smaller for stages I–II than for all stages combined, confidence intervals are narrower because probabilities are closer to 100% for this subgroup of patients. Also notice that confidence intervals do not widen with time since diagnosis, contrarily to what would be expected given that the sample size decreases with time since diagnosis. The following elements explain this phenomenon. We only consider age at diagnosis 20–69 years. Survival is high for this age range and the cancer types considered, so 10 years after diagnosis will not yet be long enough to see a clear increase in the length of the confidence intervals. And again, when survival probabilities approach 100%, the confidence intervals become smaller for a given number of observations. A similar pattern for the conditional net survival has been found in Van Ginckel et al. (2022) for female breast cancer. Calculations of the confidence intervals are further explained in Appendix A.

## 5.7 Discussion

To sum up, let us compare waiting periods obtained with our approach with results obtained according to the method proposed by Soetewey et al. (2021), which are based on the time after diagnosis when the expected present value of a standard mortgage insurance reaches the same level than the one based on XK life table. A summary is displayed in Table 5.5. We can see that waiting periods are sensibly the same for thyroid cancer patients across all methods, while they are slightly higher when estimated via the approach proposed in this paper for melanoma cancer patients. Note

Cancer site	Age at diag.	Proposed approach			Soetewey et al. (2021) All stages
		Stages I–II	Stages III–IV	All stages	
Melanoma	30	> 10	> 10	> 10	9
Melanoma	50	4	> 10	6	3
Thyroid	30	1	> 10	1	1
Thyroid	50	0	> 10	1	1
Breast	30	> 10	> 10	> 10	NA
Breast	50	7	> 10	> 10	NA

Table 5.5: Comparison of waiting periods by cancer site and age at diagnosis resulting from our approach and from Soetewey et al. (2021).

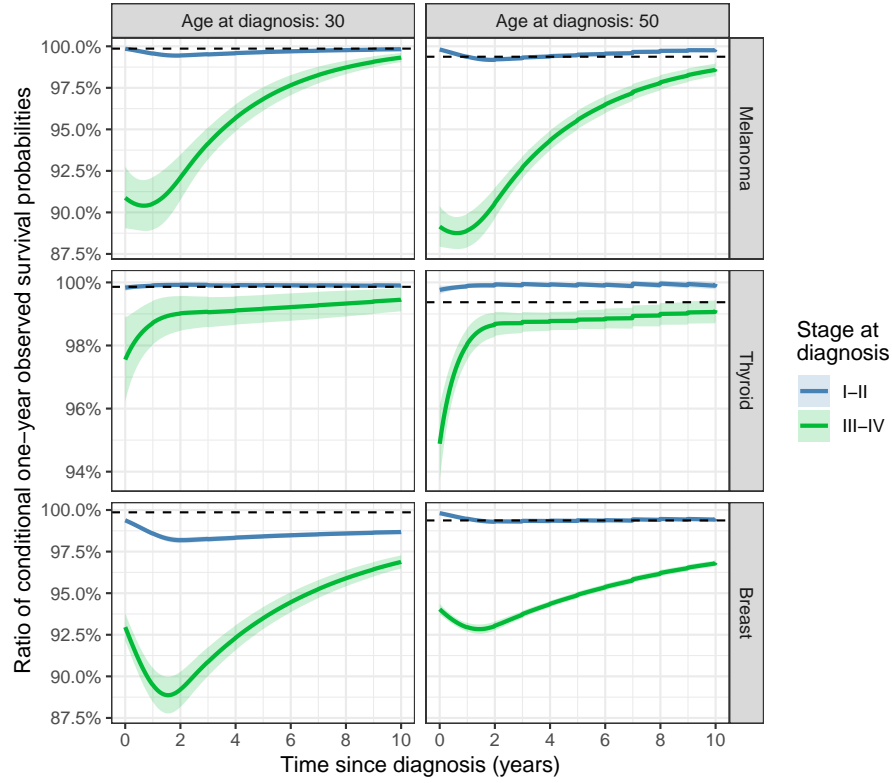


Figure 5.6: Ratio of conditional one-year survival probability obtained via the proposed approach (i.e.,  $p_{x,w}^{\text{CR}}/p_y^{\text{NIS}}$ ) by cancer site, age and stage at diagnosis. Horizontal dashed lines correspond to the additive correction  $\exp(-\gamma)$ .

that no comparison is made for breast cancer, as this cancer site was not considered in Soetewey et al. (2021).

Results in Table 5.5 are in line with the reduced waiting periods specified in the Belgian legislation. Furthermore, results are also in line with the AERAS convention (i.e., the reference grid used in France), which stipulates that the RTBF is maximum 6 years after the end of the therapeutic protocol for melanoma and thyroid cancers.

Nonetheless, an important difference is that in this paper, all waiting periods opening the RTBF are based on the time since diagnosis, rather than on the time since the end of the therapeutic protocol as currently implemented in the Belgian and French reference grids. As duration of cancer treatments are unpredictable and heterogeneous (even within the same cancer site and stage), a RTBF based on the date of diagnosis rather than based on the treatment end date will benefit both patients and insurers. Indeed, patients will know exactly when they can expect to benefit from this RTBF, and insurers will face less uncertainties (as the date of diagnosis is known and fixed, contrarily to the treatment end date which is difficult to establish and may change over time depending on the patient's health status) and less prone to debates (about, for instance, what is considered as treatment or not).

Although data used in the analyses cover a relatively long period of time (year of diagnosis ranges from 2004 to 2020) with diagnostic criteria and methods that have evolved and improved over that period, calendar time has not been included for two main reasons. First, the limited number of cases available (in particular since the focus is on young adults) prevents another division between different cohorts. Second, given that medicine and treatments progress with time, survival of cancer patients also improve with time. Thus, the resulting potential bias of omitting a cohort effect appears to be conservative, as the actual time for the patients to reach a survival comparable to that of the general population will decrease with improving treatments. In addition to that, population data are used whereas outstanding balance insurance applicants belong to the upper socio-economic class who usually have better prognosis, and individuals who contract a home or professional loan are generally in good health as individuals with poor health are unlikely to embark on such a project. These selection effects imply that analyses conducted in the present paper are conservative in many respects.

One could argue that waiting periods are expected to be shorter for patients diagnosed at stages III–IV compared to patients diagnosed at stages I–II, as we would expect when comparing patients diagnosed with pancreatic and breast cancer. The idea behind this reasoning is that the worse the prognosis, the quicker the patients die after diagnosis and thus the quicker only the survivors remain. Results of the stratification by stage show that it is not the case. The following arguments explain it. Statistical cure in the case of female breast cancer is not yet achieved within 15 years after diagnosis (except for stage I), while it is achieved for pancreatic cancer at around 5 years after diagnosis. Indeed, for female breast cancer, excess hazard is relatively constant and non negligible even after many years after diagnosis, with late recurrences occurring up to 20 years after diagnosis. On the contrary, for aggressive cancers, excess hazard is much less constant over the years after diagnosis, and in the case of pancreatic cancer it becomes negligible around 5 years after diagnosis. Given the difference in excess mortality between breast and pancreatic cancer, it is reasonable to expect waiting periods to be shorter for pancreatic than for breast cancer. Although melanoma and thyroid cancers are nowhere near as aggressive as pancreatic cancer, the trend of the excess hazard for these two cancers is closer to pancreatic than to breast cancer, that is, excess hazard is not constant over the years after diagnosis, it becomes negligible only after some years after diagnosis and late recurrences are rare. This explains the shorter waiting periods for melanoma and thyroid cancers compared to female breast cancer. The same reasoning can be applied

to the comparison of the waiting periods between stages of the tumor. One could expect that the more advanced the stage, the more quickly only the survivors remain and thus the shorter the waiting period. This holds only if statistical cure is reached at a given time after diagnosis (and in particular within 10 years after diagnosis to argue for a reduced waiting period opening the RTBF). For female breast cancer, the excess hazard for stages III–IV is higher than for stages I–II up to 15 years after diagnosis, resulting in waiting periods that are not shorter for stages III–IV compared to stages I–II.

Results obtained in the present study focus on melanoma, thyroid and female breast cancer patients for illustrative purposes. The approach developed in this paper can be applied to other cancer types or diseases. However, as just discussed, it cannot be used in case of late recurrences nor to chronic diseases to argue a shorter waiting period. For some cancers with late recurrences such as breast cancer, the waiting period resulting from our proposed approach when including patients diagnosed at all stages of the tumor is (much) longer than if it was proposed only to patients diagnosed at stages I–II. For melanoma and thyroid cancers, the waiting period resulting from our proposed approach when including patients diagnosed at all stages of the tumor is relatively similar than if it was proposed only to patients diagnosed at stages I–II. This demonstrates that, besides the fact that computing the waiting period should be done by stage for female breast cancer while it is not compulsory for melanoma and thyroid cancers, cancer is not one disease, but a family of many diverse diseases with different outcomes. Therefore, the proposed method should be applied on a case-by-case basis, that is, cancer by cancer. This is left for future research.

As mentioned in Section 5.1, the method proposed by Soetewey et al. (2021) remains actuarially sound if the length of the follow-up is long enough. It could be argued, however, that even when registry data have a sufficiently long follow-up period, the method proposed in this paper would still be preferable since a long follow-up means that some patients have been diagnosed a long time ago, and are thus not treated as well as nowadays. This argument is all the more valid the longer the follow-up time, as the longer the follow-up, the greater the potential increase in treatment efficacy between the beginning and end of the follow-up period.

The proposed approach can obviously be applied in other countries by replacing the databases by the appropriate ones. Moreover, other cancer sites and other diseases which qualify for the RTBF (e.g., HIV, some types of hepatitis and leukemia) are left for future research. This would undeniably be useful to improve the reference grids in Belgium and other countries, and ultimately, to improve access to such insurance products for other types of surviving patients.





# Conclusion

# | 6

In this thesis, we proposed new statistical methods pertaining to xxx. The objective of this work was mainly to propose new and effective tools to address some rather central questions arising in biostatistics. As one could expect, some questions that came up during this work remain unanswered. We develop upon some of them below after briefly recalling the contributions.

## 6.1 General conclusion

The current research aims to study xxx

In Chapter 2 it is shown that xxx

Chapter 2 introduces xxx

Chapter 3 extends results xxx

The key message of Chapter 4, which focuses on xxx, is that xxx

Chapter 5 of the thesis is dedicated to xxx

## 6.2 Future research

The field of xxx is currently extremely lively: the number of applications based on this theory keeps on growing steadily, while the understanding of the theory also keeps developing. Additionally to the classical definition of the problem, some strands of the literature propose variations that xxx. Still, many questions and challenges remain to be faced. We list a few potential areas for further research here.

A major gap of the thesis is the lack of xxx

In the same spirit, a possible line of further research is xxx

Another intriguing research idea would be to study xxx



# Appendix



This appendix details the calculation of the conditional one-year observed survival probabilities introduced in chapter 5.

## A.1 Conditional one-year observed survival probability

### A.1.1 Observed survival and hazard rate

From the relation between cumulative hazard for all cause death,  $\Lambda$ , follows the observed survival, OS:

$$OS(t) = \exp(-\Lambda(t)) = \exp\left(-\int_0^t \lambda(u)du\right), \quad (\text{A.1})$$

with  $\lambda(u)$  the hazard rate at time  $u$ .

Conditional one-year OS at time  $t$  can be obtained from Eq. (A.1) by integrating only over the interval  $[t, t+1]$ :

$$OS(t, t+1) = \exp\left(-\int_t^{t+1} \lambda(u)du\right). \quad (\text{A.2})$$

To calculate this integral in practice, numerical integration can be applied on a set of time values, say with a step of 0.01 year.

### A.1.2 Flexible parametric model for the hazard rate

The hazard rate as a continuous function of survival time was obtained from a flexible parametric model (fpm) using the `mexhaz` function from the R package `mexhaz` (Charvat and Belot, 2021).

### A.1.3 Predicted observed survival

The observed survival at a given time  $t$ , can be obtained from numerical integration of Eq. (A.1):

$$OS(t) = \exp\left(-\sum_{i=0}^{N_t-1} \lambda(t_i)(t_{i+1} - t_i)\right) = \exp\left(-\sum_{i=0}^{N_t-1} \lambda(t_i)\Delta t\right) = \exp(-\tilde{\Lambda}(t)), \quad (\text{A.3})$$

when the  $[0, t]$  interval is split in  $N_t$  intervals of width  $\Delta t$  ( $t_0 = 0, t_1 = \Delta t, t_2 = 2\Delta t, \dots, t_{N_t} = t$ ).

To obtain a curve of the observed survival at a set of time values (say from 0 to 10 years in steps of  $\Delta t = 0.01$  year, so 1000 data points), the vector of the corresponding cumulative hazards,  $\tilde{\Lambda}$ , calculated via numerical integration is needed:

$$OS = \exp(-\tilde{\Lambda}). \quad (\text{A.4})$$

The cumulative hazard vector can be calculated from the estimated regression coefficients via matrix multiplication. Let  $\mathbf{X}$  be the design matrix ( $N_t$  lines, each line corresponds with a time value) for the needed linear combinations of estimated regression coefficients,  $\boldsymbol{\beta}$ , at the  $\log(\lambda(t))$  scale. The estimated  $\log(\boldsymbol{\lambda})$  vector and its covariance matrix at each time points equals:

$$\log(\boldsymbol{\lambda}) = \mathbf{X}\boldsymbol{\beta} \quad (\text{A.5})$$

$$\sum_{\log(\boldsymbol{\lambda})} = \mathbf{X} \sum_{\boldsymbol{\beta}} \mathbf{X}^T. \quad (\text{A.6})$$

So:

$$\boldsymbol{\lambda} = \exp(\mathbf{X}\boldsymbol{\beta}) \quad (\text{A.7})$$

$$\sum_{\boldsymbol{\lambda}} = \mathbf{J}_{\boldsymbol{\lambda}} \sum_{\boldsymbol{\beta}} \mathbf{J}_{\boldsymbol{\lambda}}^T, \quad (\text{A.8})$$

with  $\mathbf{J}_{\boldsymbol{\lambda}}$  the Jacobian matrix  $\mathbf{J}_{\boldsymbol{\lambda}} = \text{diag}(\exp(\mathbf{X}\boldsymbol{\beta}))$ .

The cumulative hazard at all time points is easily obtained by multiplying with a upper triangular matrix  $\mathbf{T}$  (with 1's on the diagonal):

$$\tilde{\boldsymbol{\Lambda}} = \Delta t \cdot \boldsymbol{\lambda}^T \mathbf{T} \quad (\text{A.9})$$

$$\sum_{\tilde{\boldsymbol{\Lambda}}} = (\Delta t)^2 \mathbf{T}^T \sum_{\boldsymbol{\lambda}} \mathbf{T}. \quad (\text{A.10})$$

The variances of the cumulative hazards are the diagonal elements of covariance matrix, which allows to construct an asymptotic normal confidence interval (CI).

From Eq. (A.4), it follows:

$$\text{OS} = \exp(-\tilde{\boldsymbol{\Lambda}}) \quad (\text{A.11})$$

$$\sum_{\text{OS}} = \mathbf{J}_{\text{OS}} \sum_{\tilde{\boldsymbol{\Lambda}}} \mathbf{J}_{\text{OS}}^T. \quad (\text{A.12})$$

An asymptotic CI on the obtained survival can be obtained by transforming the CI on the cumulative hazard.

#### A.1.4 Predicted conditional one-year observed survival

To obtain the conditional one-year observed survival at each time point  $t_i$ , the cumulative hazard over only the next 1 year interval  $[t_i, t_i + 1]$  is needed. This can be achieved by creating an upper triangular matrix,  $\mathbf{T}_c$ , for which the number of 1's in each row is limited up to the next  $\frac{1}{\Delta t}$  columns. Take as an example  $\Delta t = 0.2$  and consider the first six lines and the first 10 columns:

$$\mathbf{T}_c = \begin{bmatrix} 1 & 1 & 1 & 1 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 1 & 1 & 1 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 1 & 1 & 1 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 1 & 1 & 1 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 1 & 1 & 1 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 1 & 1 & 1 & 1 \end{bmatrix}. \quad (\text{A.13})$$

The cumulative hazard for the conditional one-year OS becomes:

$$\tilde{\Lambda}_c = \Delta t \cdot \mathbf{T}_c \boldsymbol{\lambda} \quad (\text{A.14})$$

$$\sum_{\tilde{\Lambda}} = (\Delta t)^2 \mathbf{T}_c \sum_{\lambda} \mathbf{T}_c^T. \quad (\text{A.15})$$

The rest is similar to the observed survival in the previous subsection.



# Bibliography

- M. Amico and I. Van Keilegom. Cure models in survival analysis. *Annual Review of Statistics and Its Application*, 5:311–342, 2018.
- T. M. Andersson, P. W. Dickman, S. Eloranta, and P. C. Lambert. Estimating and modelling cure in population-based cancer studies within the framework of flexible parametric survival models. *BMC medical research methodology*, 11(1):96, 2011.
- Y. Y. Andrei, B. Asselain, et al. *Stochastic models of tumor latency and their biostatistical applications*, volume 1. World Scientific, 1996.
- J. C. Bailar III and E. M. Smith. Progress against cancer? *New England Journal of Medicine*, 314(19):1226–1232, 1986.
- C. M. Balch, S.-J. Soong, J. E. Gershenwald, J. F. Thompson, D. S. Reintgen, N. Cascinelli, M. Urist, K. M. McMasters, M. I. Ross, J. M. Kirkwood, et al. Prognostic factors analysis of 17,600 melanoma patients: validation of the american joint committee on cancer melanoma staging system. *Journal of clinical oncology*, 19(16):3622–3634, 2001.
- A. Belot, A. Ndiaye, M.-A. Luque-Fernandez, D.-K. Kipourou, C. Maringe, F. J. Rubio, and B. Rachet. Summarizing and communicating on survival data according to the audience: a tutorial on different measures illustrated with population-based cancer registry data. *Clinical epidemiology*, 11:53–65, 2019.
- C. L. Bennett, P. D. Weinberg, and J. J. Lieberman. Cancer insurance policies in japan and the united states. *Western journal of medicine*, 168(1):17, 1998.
- J. Berkson and R. P. Gage. Calculation of survival rates for cancer. *Proceedings of the staff meetings. Mayo Clinic*, 25(11):270–286, May 1950. ISSN 0092-699X. URL <http://europepmc.org/abstract/MED/15417650>.
- J. Berkson and R. P. Gage. Survival curve for cancer patients following treatment. *Journal of the American Statistical Association*, 47(259):501–515, 1952.
- F. Berrino, R. Capocaccia, J. Estève, G. Gatta, T. Hakulinen, A. Micheli, M. Sant, and A. Verdecchia. Survival of cancer patients in europe: the eurocare-2 study. In *Survival of cancer patients in Europe: The EURO CARE-2 study*, volume 151, pages 1–CD. IARC Scientific Publication, 1999.
- J. W. Boag. Maximum likelihood estimates of the proportion of patients cured by cancer therapy. *Journal of the Royal Statistical Society. Series B (Methodological)*, 11(1):15–53, 1949.
- P. Bolard, C. Quantin, J. Esteve, J. Faivre, and M. Abrahamowicz. Modelling time-dependent hazard ratios in relative survival: application to colon cancer. *Journal of clinical epidemiology*, 54(10):986–996, 2001.
- O. Boussari, G. Romain, L. Remontet, N. Bossard, M. Mounier, A.-M. Bouvier, C. Biquet, M. Colonna, and V. Jooste. A new approach to estimate time-to-cure from cancer registries data. *Cancer epidemiology*, 53:72–80, 2018.
- J. Buckley. Additive and multiplicative models for relative survival rates. *Biometrics*, 40:51–62, 1984.
- Bureau du suivi de la tarification. Rapport sur l’activité 2017. [https://www.bureaudusuivi.be/images/docs/RapportAnnuel\\_2017.pdf](https://www.bureaudusuivi.be/images/docs/RapportAnnuel_2017.pdf), 2018.

- A. E. Chang, P. A. Ganz, D. F. Hayes, T. Kinsella, H. I. Pass, J. H. Schiller, R. M. Stone, and V. Strecher. *Oncology: an evidence-based approach*. Springer Science & Business Media, 2007.
- H. Charvat and A. Belot. mexhaz: An R package for fitting flexible hazard-based regression models for overall and excess mortality with a random effect. *Journal of Statistical Software*, 98(14):1–36, 2021. doi: 10.18637/jss.v098.i14.
- M. Chauvenet, C. Lepage, V. Jooste, V. Cottet, J. Faivre, and A.-M. Bouvier. Prevalence of patients with colorectal cancer requiring follow-up or active treatment. *European Journal of Cancer*, 45(8):1460–1465, 2009.
- M.-H. Chen, J. G. Ibrahim, and D. Sinha. A new bayesian model for survival data with a surviving fraction. *Journal of the American Statistical Association*, 94(447):909–919, 1999.
- M. Clements and X.-R. Liu. *rstpm2: Smooth Survival Models, Including Generalized Survival Models*, 2019. URL <https://CRAN.R-project.org/package=rstpm2>. R package version 1.5.1.
- I. Clerc-Urmès and M. Grzebyk. *flexrsurv: Flexible Relative Survival Analysis*, 2023. URL <https://CRAN.R-project.org/package=flexrsurv>. R package version 2.0.17.
- I. Clerc-Urmès, M. Grzebyk, G. Hédélec, and CENSUR working survival group. *flexrsurv: An R package for relative survival analysis*, 2020. URL <https://CRAN.R-project.org/package=flexrsurv>. R package version 1.4.5.
- W. Cleveland, E. Grosse, M. Shyu, E. Grosse, M. Shyu, and M. Shyu. A package of c and fortran routines for fitting local regression models. *Statistical methods, edited by: Chambers, JM, S. Chapman and Hall Ltd., London, UK*, 1992.
- W. S. Cleveland and E. Grosse. Computational methods for local regression. *Statistics and computing*, 1(1):47–62, 1991.
- W. S. Cleveland, S. J. Devlin, and E. Grosse. Regression by local fitting: methods, properties, and computational algorithms. *Journal of econometrics*, 37(1):87–114, 1988.
- M. P. Coleman, P. Babb, P. Damiecki, P. Grosclaude, S. Honjo, J. Jones, G. Knerer, A. Pitard, M. Quinn, A. Sloggett, et al. *Cancer survival trends in England and Wales, 1971-1995: deprivation and NHS region*. Stationery Office Books, 1999.
- L. Dal Maso, S. Guzzinati, C. Buzzoni, R. Capocaccia, D. Serraino, A. Caldarella, A. Dei Tos, F. Falcini, M. Autelitano, G. Masanotti, et al. Long-term survival, prevalence, and cure of cancer: a population-based estimation for 818 902 italian patients and 26 cancer types. *Annals of oncology*, 25(11):2251–2260, 2014.
- C. Danieli, L. Remontet, N. Bossard, L. Roche, and A. Belot. Estimating net survival: the importance of allowing for informative censoring. *Statistics in medicine*, 31(8):775–786, 2012.
- J. Dębicka, A. Marciniuk, and B. Zmyślona. Combination reverse annuity contract and critical illness insurance. In *18th AMSE, Applications of Mathematics and Statistics in Economics, Conference Proceedings*, pages 2–6, 2015.
- M. Denuit and C. Legrand. Risk classification in life and health insurance: extension to continuous covariates. *European Actuarial Journal*, 8:245–255, 2018.
- M. Denuit, N. Lucas, and E. Pitacco. Pricing and reserving in ltc insurance. *Actuarial aspects of long term care*, pages 129–158, 2019.
- P. W. Dickman, A. Sloggett, M. Hills, and T. Hakulinen. Regression models for relative survival. *Statistics in medicine*, 23(1):51–64, 2004.
- D. C. Dickson, M. Hardy, M. R. Hardy, and H. R. Waters. *Actuarial mathematics for life contingent risks*. Cambridge University Press, 2013.
- E. O. Dodd, G. Streftaris, H. R. Waters, and A. D. Stott. The effect of model uncertainty on the pricing of critical illness insurance. *Annals of Actuarial Science*, 9(1):108–133, 2015.



- P. H. Eilers and B. D. Marx. Flexible smoothing with b-splines and penalties. *Statistical science*, 11(2):89–102, 1996.
- J. E. Enstrom and D. F. Austin. Interpreting cancer survival rates. *Science*, 195(4281):847–851, 1977.
- J. Esteve, E. Benhamou, M. Croasdale, and L. Raymond. Relative survival and the estimation of net survival: elements for further discussion. *Statistics in medicine*, 9(5):529–538, 1990.
- J. Esteve, E. Benhamou, L. Raymond, et al. Statistical methods in cancer research. volume iv. descriptive epidemiology. *IARC Sci publ*, 128(1):302, 1994.
- M. Fauvernier, L. Remontet, Z. Uhry, N. Bossard, and L. Roche. survpen: an r package for hazard and excess hazard modelling with multidimensional penalized splines. *Journal of Open Source Software*, 4(40):1434, 2019a. doi: 10.21105/joss.01434.
- M. Fauvernier, L. Roche, Z. Uhry, L. Tron, N. Bossard, L. Remontet, and C. in the Estimation of Net Survival Working Survival Group. Multi-dimensional penalized hazard model with continuous covariates: applications for studying trends and social inequalities in cancer survival. *Journal of the Royal Statistical Society: Series C (Applied Statistics)*, 2019b.
- R. B. Geskus. *Data analysis with competing risks and intermediate states*, volume 82. CRC Press, 2015.
- R. Giorgi, J. Payan, and J. Gouvernet. Rsurv: a function to perform relative survival analysis with s-plus or r. *Computer methods and programs in biomedicine*, 78(2):175–178, 2005.
- S. Haberman and A. Renshaw. Generalised linear models and excess mortality from peptic ulcers. *Insurance: Mathematics and Economics*, 9(1):21–32, 1990.
- T. Hakulinen and L. Tenkanen. Regression analysis of relative survival rates. *Journal of the Royal Statistical Society: Series C (Applied Statistics)*, 36(3):309–317, 1987.
- M. J. Hendriks, E. Harju, K. Roser, M. Ienca, and G. Michel. The long shadow of childhood cancer: a qualitative study on insurance hardship among survivors of childhood cancer. *BMC Health Services Research*, 21(1):1–12, 2021.
- L. H. Jakobsen, T. M.-L. Andersson, J. L. Bickler, L. Ø. Poulsen, M. T. Severinsen, T. C. El-Galaly, and M. Bøgsted. On estimating the time to statistical cure. *BMC Medical Research Methodology*, 20:1–13, 2020.
- V. Jooste, P. Grosclaude, L. Remontet, G. Launoy, I. Baldi, F. Molinié, P. Arveux, N. Bossard, A.-M. Bouvier, M. Colonna, et al. Unbiased estimates of long-term net survival of solid cancers in france. *International journal of cancer*, 132(10):2370–2377, 2013.
- E. L. Kaplan and P. Meier. Nonparametric estimation from incomplete observations. *Journal of the American statistical association*, 53(282):457–481, 1958.
- P. C. Lambert, J. R. Thompson, C. L. Weston, and P. W. Dickman. Estimating and modeling the cure fraction in population-based cancer survival analysis. *Biostatistics*, 8(3):576–594, 2006.
- A. Latouche, A. Allignol, J. Beyersmann, M. Labopin, and J. P. Fine. A competing risks analysis should report results on all cause-specific hazards and cumulative incidence functions. *Journal of clinical epidemiology*, 66(6):648–653, 2013.
- M. Lawler, F. De Lorenzo, P. Lagergren, F. S. Mennini, S. Narbutas, G. Scocca, F. Meunier, and E. A. of Cancer Sciences. Challenges and solutions to embed cancer survivorship research and innovation within the eu cancer mission. *Molecular Oncology*, 15(7):1750–1758, 2021.
- C. Legrand and A. Bertrand. Cure models in oncology clinical trials. *Textbook of Clinical Trials in Oncology: A Statistical Perspective*, 1:465–492, 2019.
- J. Lemaire, K. Subramanian, K. Armstrong, and D. A. Asch. Pricing term insurance in the presence of a family history of breast or ovarian cancer. *North American Actuarial Journal*, 4(2):75–87, 2000.

- P. Lenner. The excess mortality rate: a useful concept in cancer epidemiology. *Acta Oncologica*, 29(5):573–576, 1990.
- X.-R. Liu, Y. Pawitan, and M. S. Clements. Generalized survival models for correlated time-to-event data. *Statistics in medicine*, 36(29):4743–4762, 2017.
- X.-R. Liu, Y. Pawitan, and M. Clements. Parametric and penalized generalized survival models. *Statistical methods in medical research*, 27(5):1531–1546, 2018.
- R. A. Maller and X. Zhou. *Survival analysis with long-term survivors*. John Wiley & Sons, 1996.
- M. Massart. A long-term survivor’s perspective on supportive policy for a better access to insurance, loan and mortgage. *Journal of cancer policy*, 15:70–71, 2018.
- M. Mounier. *Apport des méthodes de survie nette dans le pronostic des lymphomes malins non hodgkiniens en population générale*. PhD thesis, Université Claude Bernard-Lyon I, 2015.
- R. B. Nielsen and R. N. Mayer. Why do people buy cancer insurance? an exploratory study. *Advancing the Consumer Interest*, pages 16–22, 2000.
- H. Oksanen. *Modelling the survival of prostate cancer patients*. University of Tampere, 1998.
- M. Othus, B. Barlogie, M. L. LeBlanc, and J. J. Crowley. Cure models as a useful statistical tool for analyzing survival. *Clinical Cancer Research*, 18(14):3731–3736, 2012.
- K. Pavlič and M. Pohar Perme. Using pseudo-observations for estimation in relative survival. *Biostatistics*, 20(3):384–399, 2019.
- C. Percy, E. Stanek 3rd, and L. Gloeckler. Accuracy of cancer death certificates and its effect on cancer mortality statistics. *American Journal of Public Health*, 71(3):242–250, 1981.
- M. P. Perme and K. Pavlič. Nonparametric relative survival analysis with the R package relsurv. *Journal of Statistical Software*, 87(8):1–27, 2018. doi: 10.18637/jss.v087.i08.
- M. P. Perme, J. Stare, and J. Estève. On estimation in relative survival. *Biometrics*, 68(1):113–120, 2012.
- M. P. Perme, J. Estève, and B. Rachet. Analysing population-based cancer survival—settling the controversies. *BMC cancer*, 16(1):1–8, 2016.
- E. Pitacco. Health insurance. *Basic Actuarial Models*, Cham, Switzerland: Springer Verlag, 2014.
- M. Pohar and J. Stare. Relative survival analysis in r. *Computer methods and programs in biomedicine*, 81(3):272–278, 2006.
- C. Quantin, M. Abrahamowicz, T. Moreau, G. Bartlett, T. MacKenzie, M. Adnane Tazi, L. Lalonde, and J. Faivre. Variation over time of the effects of prognostic factors in a population-based study of colon cancer: comparison of statistical models. *American journal of epidemiology*, 150(11):1188–1200, 1999.
- R Core Team. *R: A Language and Environment for Statistical Computing*. R Foundation for Statistical Computing, Vienna, Austria, 2017. URL <https://www.R-project.org>.
- L. Remontet, Z. Uhry, N. Bossard, J. Iwaz, A. Belot, C. Danieli, H. Charvat, L. Roche, and C. W. S. Group. Flexible and structured survival model for a simultaneous estimation of non-linear and non-proportional effects and complex interactions between continuous variables: Performance of this multidimensional penalized spline approach in net survival trend analysis. *Statistical methods in medical research*, 28(8):2368–2384, 2019.
- A. E. Renshaw. Modelling excess mortality using glim. *Journal of the Institute of Actuaries*, 115(2):299–315, 1988.
- L. Ries, M. Eisner, C. Kosary, B. Hankey, B. Miller, L. Clegg, and B. Edwards. Seer cancer statistics review, 1973–1999. bethesda, md: National cancer institute. Also available from: URL: <http://seer.cancer.gov/csr/1973–1999>, 2002.

- R. Schaffar, B. Rachet, A. Belot, and L. M. Woods. Estimation of net survival for cancer patients: relative survival setting more robust to some assumption violations than cause-specific setting, a sensitivity analysis on empirical data. *European Journal of Cancer*, 72:78–83, 2017.
- G. Schvartsman, P. Taranto, I. C. Glitza, S. S. Agarwala, M. B. Atkins, and A. C. Buzaid. Management of metastatic cutaneous melanoma: updates in clinical practice. *Therapeutic Advances in Medical Oncology*, 11:1758835919851663, 2019.
- G. Scocca and F. Meunier. A right to be forgotten for cancer survivors: a legal development expected to reflect the medical progress in the fight against cancer. *Journal of Cancer Policy*, 25:100246, 2020.
- G. Scocca and F. Meunier. Towards an EU legislation on the right to be forgotten to access to financial services for cancer survivors. *European Journal of Cancer*, 162: 133–137, 2022.
- K. Shang. *Individual Cancer Mortality Prediction*. Fundaciòn Mapfre, sep 2019. ISBN 8498446651. URL <https://www.xarg.org/ref/a/8498446651/>.
- G. Silversmit, D. Jegou, E. Vaes, E. Van Hoof, E. Goetghebeur, and L. Van Eycken. Cure of cancer for seven cancer sites in the Flemish region. *International journal of cancer*, 140(5):1102–1110, 2017.
- A. Soetewey, C. Legrand, M. Denuit, and G. Silversmit. Waiting period from diagnosis for mortgage insurance issued to cancer survivors. *European Actuarial Journal*, 11 (1):135–160, 2021.
- A. Soetewey, C. Legrand, M. Denuit, and G. Silversmit. Semi-markov modeling for cancer insurance. *European Actuarial Journal*, 12(2):813–837, 2022.
- Terry M. Therneau and Patricia M. Grambsch. *Modeling Survival Data: Extending the Cox Model*. Springer, New York, 2000. ISBN 0-387-98784-3.
- The Belgian Cancer Registry. Cancer survival in belgium. <https://kankerregister.org/media/docs/publications/CancerSurvivalinBelgium.PDF>, 2012.
- A. Tsodikov, J. Ibrahim, and A. Yakovlev. Estimating cure rates from survival data: an alternative to two-component mixture models. *Journal of the American Statistical Association*, 98(464):1063–1078, 2003.
- A. Van Ginckel, G. Silversmit, B. Van Gool, N. Van Damme, and P. Jonckheer. The right to be forgotten in breast cancer: new propositions. *Belgian Health Care Knowledge Centre (KCE)*, KCE Reports 351, 2022. doi: 10.57598/R351C.
- J. C. Yue, H.-C. Wang, Y.-Y. Leong, and W.-P. Su. Using taiwan national health insurance database to model cancer incidence and mortality rates. *Insurance: Mathematics and Economics*, 78:316–324, 2018.
- Y. Zhan, X.-R. Liu, C. A. Reynolds, N. L. Pedersen, S. Hägg, and M. S. Clements. Leukocyte telomere length and all-cause mortality: a between-within twin study with time-dependent effects using generalized survival models. *American Journal of Epidemiology*, 187(10):2186–2191, 2018.