Commitment and Lapse Behavior in Long-Term Insurance: A Case Study*

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Abstract

This paper presents a case study of a portfolio of individual long-term insurance contracts sold by a Spanish mutual company. We describe the risk levels, the rating structure and the implied cross-subsidies on a portfolio of policies providing health, life and long-term care insurance. We show evidence of reclassification risk through the history of disability spells. We also analyze the lapse behavior and seek to provide a rationale for the portfolio's dynamics. We discuss the lack of commitment from the policyholders (lapses) and from the mutual company (which took a run-off decision). Finally, we draw conclusions regarding the design of such contracts.

1 Introduction

This paper presents a case study of a portfolio of individual long-term insurance contracts, sold by a Spanish mutual company. The portfolio has been set in a run-off position - i.e. has been closed to new business - since 1997. Our reasons

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for undertaking such an analysis are threefold: a) empirical studies of complex contracts such as the one studied here are extremely rare; b) commitment and lapse behavior can be studied here using a data set that includes information on a portfolio for a period extending over more than two decades; and, c) conclusions can be drawn regarding the consequences of closing a portfolio to new business, while keeping the existing contracts in that portfolio. The contract comprises a bundle of three coverages for health, long-term care (hereinafter LTC) and life insurance. The life coverage combines term and whole life insurance. The health coverage is unfunded (i.e., current premiums finance current benefits, and reserves are set only for claims incurred in the current period). By contrast, the life and LTC policies exhibit a more complex funding scheme, which is discussed in more detail below (see Section 4). As is usual with long-term contracts, there is a one-sided commitment in terms of loyalty. So while the policyholders can leave the mutual company, the company cannot cancel the contract. Consequently, the policyholder is insured against reclassification risk, given that experience rating is also forbidden. However, the insurer is not committed to a long-term premium scheme, and the average premium level follows the average loss trend in an unfunded setting. If the premium-benefit ratio only depends on calendar time, the insurance company follows a "community rating" strategy.

Risks related to disability and death increase with age, but are also subject to marked cohort effects. Due to mortality improvements, insurance companies benefit from these effects as regards death benefit insurance (whether term life or whole life insurance). However, as a result of aging, long-term care risk increases with calendar time. An insurance company's natural hedge against uncertainty in the Knightian sense is not to commit itself to a long-term premium scheme. In our study, short-term risks increase with age much more rapidly than do the corresponding premiums. Besides, gender is not taken into account. These characteristics entail strong cross-subsidies between genders and generations. The difference between the insurance premium and the corresponding risk level is a subsidy in an unfunded setting, and a savings in a funded scheme. For sake of simplicity, we will use the first terminology in this paper, although two of the three components in the bundle do incorporate some funding. Finally, there is a surrender value associated with the death benefit component, but none associated with the LTC coverage.

The article is organized as follows. Section 2 reviews the literature on long-term insurance. Section 3 describes the insurance contract, and the portfolio analyzed in the empirical study. The database contains 150,000 individual insurance contracts with a history of up to thirty years. Related variables include the purchase date, the history of disability spells (with their respective initiation and termination dates), and the date of entry in an LTC spell for permanently disabled policyholders. We have the cancellation date of closed contracts and the related cause (death or lapse at the initiative of the policyholder). This section also provides an outline of the economic framework, and more specifically the evolution in public and private health insurance in Spain. The three risks covered by the insurance bundle are described in Section 4. Our study focuses on health risk, given that life risk is well known and LTC risk for this portfolio is analyzed in a previous paper (Guillén and Pinquet, 2008). A further reason is that the history of disability spells is key in the learning on a policyholder's health status, and hence may help explain lapsation behavior. We show evidence of reclassification risk through the history of disability spells. We link the rating structure of life and health policies, depending on the nature of funding (fully funded, unfunded with or without cross-subsidies between age classes) to three variables (calendar time, seniority in the portfolio, age at inception of the contract). We then analyze the rating structure of the mutual company. We find that the company follows a strict "community rating" strategy for its health coverage, and that young policyholders heavily subsidize the older policyholders for all the guarantees.

Losses incurred by lapses can be high for front-loaded contracts without surrender value. This is the case for most LTC policies, and motivates an analysis of lapse behavior.[1] The average lapse rate of LTC contracts stands at 7% in the US (Society of Actuaries, 2002). Lapse rates can be even higher if policyholders are enrolled in the contract, as in the "ElderShield" program in Singapore.[2] In a given cohort of LTC insurance purchasers, a majority will thus end their life cycle without coverage, while the probability of entering an LTC spell before dying increases with age. We provide empirical evidence on the lapse behavior in Section 5, and seek to provide a rationale for the results. We find that policyholders who cancel their contract have good health histories compared to those of their peers, and that the lapse rate decreases with age, with a local peak at 65 years. We argue that lapsation of young policyholders as well as that of elderly policyholders at retirement is partly due to a misunderstanding

[1] For instance, fewer than 3% of the US policies analyzed by Brown and Finkelstein (2007) provide any benefits once a policy lapses.

[2] In 2002, a mandatory LTC coverage was introduced for Singaporeans aged between 40 and 70. The opt-out option was retained by just 15% of this population. However, the lapse rate stood at 38% during the first year. It subsequently fell, but remained at 14% in 2006.

of the contract. We also discuss the fact that the portfolio avoids the "death spiral" that might have been expected after the run-off decision taken in 1997, caused by the continuous departure of the youngest policyholders. In Section 6, we summarize our results and discuss the design of long-term contracts.

2 A review of the literature on long-term insurance

Issues related to commitment, cross-subsidies between periods and lapse behavior in long-term insurance contracts have already been addressed extensively in the economic literature. Cross-subsidies between the periods of a contract are termed as either "lowballing" or "highballing", depending on whether the first periods are subsidized by the following ones, or the contrary. Some studies adopt alternative terminology and speak of "back-loading" and "front-loading" respectively. The contracts analyzed in this article are of the "front-loading" type. Young policyholders, although they pay less than the older ones, heavily subsidize them as will be shown in Section 4. "Lowballing" in insurance contracts may occur when the insurer extracts a rent from the policyholder based on its use of private information (Kunreuther and Pauly, 1985).[3] In our database, such information can be obtained from the history of disability spells, but experience rating is forbidden in the long-term contract analyzed here. Fluet, Schlesinger and Fei (2009) discuss multiperiod contracts with an opting-out or opting-in opportunity, the price of which must be paid in advance. These contracts include front-loading and are of interest when the motivation to insure varies over time. As argued by Pauly, Kunreuther and Hirth (1995), risks that evolve unpredictably (such as those related to health and life) are more likely to be subject to "highballing" and to guaranteed renewability of contracts. Dionne and Doherty (1994) present a "highballing" two-period model with adverse selection, unilateral switching (i.e. one-sided commitment), and renegotiation. If the insurer commits to a premium scheme in the second period with experience rating, low risks can choose this type of contract rather than a short-term one. Closer to our setting is the study undertaken by Hendel and Lizzeri (2003) who empirically analyze term life insurance linked to a model with symmetric learning, one-sided commitment and buyer heterogeneity in the cost of front-loading.

[3] The model is derived in a no-commitment setting, with myopic consumers (i.e. those who take decisions based on the current contract).

Their model predicts that lapse rates decrease with the level of front-loading, as does the average risk in the second period, because lapses affect the lowest risk levels reached after reclassification.[4] This result is confirmed empirically by the average premiums observed in the USA for three types of term life insurance contracts (either with yearly updated premiums, or with levels in premiums and front-loading, or with state contingent prices).

Lapses in long-term insurance contracts strongly influence the *ex post* balance of the coverage, as the surrender value, if it exists at all, is often very low compared to a retrospective actuarial value.[5] Based on a statistical study of American long-term care contracts, Brown and Finkelstein (2007, 2009) derive a virtual loading factor equal to 0.18 when the policies are held until death, whereas the actual loading factor (i.e. when lapses are accounted for) is equal to 0.51[6]. Besides, the losses incurred by lapsing deter policyholders from switching to another contract, and a locked-in customer faces higher intermediation costs. Loading factors also reflect increasing returns to scale with respect to the size of the insured groups. As for health insurance, Diamond (1992) mentions a 40 percent loading factor for groups of five or fewer, and a 5.5 percent loading factor for groups of 10,000 or more. In our portfolio of individual policies, the loading factor is, unsurprisingly, closer to the first figure quoted by Diamond.

Policyholders may lapse for reasons related to risk: Finkelstein, McGarry and Sufi (2005) estimate that the participants on the Health and Retirement Survey who allowed their LTC policy to lapse are later 35 percent less risky than their peers with respect to LTC risk at a five-year horizon.[7] This dynamic risk-based selection effect lowers the efficiency of insurance against reclassification risk. Finkelstein, McGarry and Sufi (2005) also report that individuals who lapse are substantially poorer and less educated than individuals who do not lapse. This evidence on wealth suggests that liquidity constraints are another

[4]In a two-period model, the surplus generated by front-loading in optimal contracts subsidizes the policyholders that reach high risk levels after reclassification. Policyholders in the lowest risk levels after reclassification are rated at the spot price so as to deter them from lapsing.

[5]An exception is the whole life insurance coverage. Besides, these contracts are assignable, which allows secondary markets to be created. The associated arrangements are termed "life settlements" and "viatical settlements". See Doherty, Singer (2002), and Daily, Hendel and Lizzeri (2008) for the pros and cons of secondary life insurance markets.

[6]Derivations are performed at the horizon of the life cycle and in a funded setting.

[7]This difference could also be explained by a moral hazard effect. However, a comparison of LTC risk for those who lapse to a new contract versus those who lapse to no contract leads the authors to reject this effect.

factor contributing to lapsation. Moreover, this cause of lapsation has unpleasant redistribution effects, as poor policyholders subsidize richer ones. To the best of our knowledge, comprehensive empirical analyses of the causes of lapsation in long-term insurance contracts remain unavailable.

Long-term care insurance can also be sold in tandem with an annuity product. In a recent special issue of this journal, Webb (2009) analyzes a bundle of LTC and deferred annuity coverages. In a model with two risk aversion levels and a link between risk aversion and health status, an equilibrium is possible in a single market for bundled contracts. The condition is that selection effects for the two types of risk averse individuals work in opposite directions in the markets for stand-alone contracts. Murtaugh et al. (2001) examine the implications of the positive correlation of mortality and disability for the benefits of combining an immediate income annuity with long-term care disability coverage at retirement ages. They show that combining the two products could reduce the cost of both coverages and make them available to a greater number of people by reducing adverse selection in the income annuity and minimizing the need for medical underwriting for disability coverage. The complementarity between LTC insurance and annuities also depends on the other assets of the household. Davidoff (2009) shows that if consumers typically liquidate home equity only in the event of illness or very old age, then LTC insurance and annuities become less attractive and may become substitutes rather than complements.[8]

Finally, let us mention other studies of long-term insurance recently published in this journal. McShane and Cox (2009) analyze the participation of US insurers in the LTC market. They find that participation and volume decisions are made independently. Smoluk (2009) examines the relationship between long-term disability claim rates and the consumption-to-wealth ratio. Using cointegration analysis, he finds a decreasing link between these two variables in a long-run equilibrium setting.

3 A Spanish portfolio of long-term insurance contracts

We analyze a sample of 150,000 individual insurance contracts drawn at random at a fifty percent rate from a mutual insurance company in July 2007. The con-

[8]The reason is that the marginal utility of wealth drops when an otherwise illiquid home is sold, an event correlated with the payouts of both annuities and LTC insurance.

tracts comprise a bundle of three policies: death benefit insurance, health coverage and a long-term care component. The products could be bought separately, but the bundle was promoted by the company so as to minimize underwriting costs. No age constraint is applied to the benefits, and all coverages extend into the policyholders' whole lifespan regardless of their employment status (employed, retired or other non-active status). As is typical for long-term contracts, there is a one-sided commitment in terms of loyalty. The policyholders can leave the mutual company, but the company cannot cancel the contract. Hence, attrition is attributable solely to cancellation at the initiative of the policyholder, or to death. Due to modifications in the computing environment in 1992, the contracts that were closed before that date were removed from the database. Although data are available from 1975, we have restricted our analysis from 1993 to 2006 to contracts that are either working or cancelled after 1992. The variables related to the history are the purchase date, the start and end dates of disability spells, and the date of entry in an LTC spell for permanently disabled policyholders. We also record the cancellation date of closed contracts and the cause (death or lapse at the initiative of the policyholder). Other available variables include gender, date of birth, and the values for the last premiums and benefits on each policy. We have no information regarding reserves, which leads us to question the nature of the funding. In the time dimension, these data are richer than panel databases such as HRS and SHARE obtained from successive surveys, but we have far fewer variables.[9] Premiums are paid on a monthly basis, and their average values are 5 Euros for the death benefit, 6.5 for health coverage and 2.3 for the long-term care component. Disability benefits are paid daily, whereas benefits related to LTC spells are paid monthly. The average death benefit for working contracts is equal to 1,600 Euros, and average monthly benefits for disability and LTC are 75 and 115 Euros respectively. These low premium and benefit levels indicate that the product is merely additional to the public coverage. They also adapt to the financial means of policyholders who are young at the inception of the contract (see Table 1). Benefits are indexed to an inflation rate. The ratio premium-benefit varies over time without precommitment (see Section 4 for results). Benefits and premiums cannot be modified on an individual basis. This applies to both reasons of risk (experience

[9] For instance, we do not have any financial data for the household. Besides, we cannot address issues related to insurance demand from a database of policyholders. We do not know either whether lapses are followed or not by switching to another contract. HRS and SHARE are the acronyms for the Health and Retirement Survey, and the Survey of Health, Ageing and Retirement in Europe, respectively.

rating) and modifications of the coverage level. The insurance package can only be cancelled completely, and only the death benefit component has a surrender value. The LTC coverage does not include any nonforfeiture clause (i.e. a right to receive reduced benefits if the policyholder lapses the contract beyond a given seniority threshold).

Health coverage works when the policyholder temporarily requires medical treatment and cannot perform any daily activities. The state of disability is assessed by doctors appointed by the insurance company on the basis of standard medical and physical tests. There is no connection between the compensation decision made by the company and those taken by the Public Health and Social Security agencies, concerning workers compensation subsidies or disability pension. Long-term care covers individuals with a severe dependence level and who are not able to perform daily life activities without the assistance of another person. Eligibility conditions are particularly strict. The contract defines eligibility for such coverage as a permanent and irreversible loss of the capacity to function autonomously due to: irreversible psychotic disorder, hemiplegia, paraplegia, severe Parkinson disorder, aphasia or Wenicke disorder, or dementia due to cerebral malfunction. In addition, following the company's traditional practices, blindness or loss of two arms or legs are sufficient conditions to grant compensation. Finally, death benefit is not restricted in any way by the cause or place of death.

Let us give some details on the origins of this insurance contract. At the inception of the product, a group of insurance agents decided to set up a fund that could compensate women facing the untimely death of their husband. Later, death benefit coverage was extended in order to meet the needs of a policyholder throughout the life cycle, as seen by the company agents. People underwriting the product were not called "insured" but "affiliated", a term usually retained for those enrolled on a social security scheme. This product became popular among exclusively male workers. Women did not become affiliated until the 70s, given that in Spain they did not participate in the job market before that date. In this decade, the Spanish government instituted the so-called "Development Plans" (Planes de Desarrollo) that fostered social and economic change and allowed women to enter the job market for the first time. A proper public health and welfare system was also created in the late 70s. Initially, this system offered health coverage, but later it included disability pensions, unemployment subsidies and more recently long-term care. Before the creation of the public welfare system, mutual companies offered a form of private insurance to a growing working class. As an example, the mutual company analyzed in this article had roughly 60,000 affiliates in 1960. This number rose to 170,000 in 1975 and to 250,000 in 1984. The product was originally distributed in Catalonia by the mutual company's agents. During the 80s, the product was also sold in other Spanish regions as the company expanded. However, the company stopped selling the product in 1997, and from that time on, the coverages were sold separately.

Table 1 shows basic statistics regarding the portfolio dynamics for each calendar year: the average age of policyholders (x), the average age at entry, at cancellation and at death (x_e, x_c, and x_d), as well as the corresponding rates (r_e, r_c, and r_d, as a percentage of the portfolio size). We observe a steady aging of the portfolio, but the rate and causes of this aging vary with time. If the variables included in the table were defined in continuous time, the time derivative of the average age would be equal to

$$x^{'} = 1 + r_e\,(x_e - x) - r_c\,(x_c - x) - r_d\,(x_d - x)\,.$$

During the first ten years, the age at cancellation is lower on average than the age of policyholders, which contributes to the aging of the portfolio. The lapse rate increases when the portfolio is set in a run-off position, and falls thereafter. The most striking feature of Table 1 is the evolution of age at cancellation once the portfolio is closed to new business. The average first increases steadily, then dramatically at the end of the period, which suggests a modification in the motivations for lapsation.

Table 1
Descriptive statistics for the portfolio

Year	x	x_e	x_c	x_d	r_e	r_c	r_d
1993	45.4	30.2	37.3	68.7	3	4.5	0.6
1994	46	30.9	38.6	69.4	5.4	5.4	0.8
1995	46.4	33.7	37.2	70.4	5.4	6.4	0.8
1996	47.2	37.2	38.8	71.3	7.3	6.7	0.8
1997	48.2	35.7	39.1	71.9	2.3	10.4	0.9
1998	49.9		40.8	72.9		10.1	0.9
1999	51.6		42.7	73.8		8.9	1.1
2000	53.1		46.2	74.5		8.0	1.1
2001	54.2		47.4	74.7		7.3	1.2
2002	55.4		48.4	76.3		7.4	1.4
2003	56.6		51.7	76.9		5.9	1.5
2004	57.2		67.3	77.2		2.4	1.6
2005	57.6		66.3	78.3		2.4	1.7
2006	58.1		64.8	79.1		2.5	1.6

Average age of policyholders, average age at entry, at cancellation and at death: (x, x_e, x_c, and x_d). Entry, cancellation and death rates: r_e, r_c, and r_d (expressed as percentage of the portfolio size)

4 The three coverages: their rating structure, risk levels, and implied cross-subsidies

4.1 The rating structure

The rating structure of the three components of the insurance package is described in Table 2, based on an analysis of the premium-benefit ratio. We estimate a linear model where the dependent variable is the logarithm of the premium-benefit ratio and where the covariates are gender, age at the inception of the contract, the seniority of the policyholder and binary variables related to the year with available premiums. The first objective is to question the nature of funding from these ratios, as we do not have any information regarding the reserves. Life and health risks increase with age, with a stronger age effect for life and LTC risks than for health risk. In this setting, three covariates retained in the regression (age at the inception of the contract, seniority in the portfolio, and calendar time) have expected effects on the rating structure which depend on the type of funding.

- For a fully funded coverage, the main effect is expected from the age at the inception of the contract. Indeed, the premium-benefit ratio is derived from an expected balance between discounted premiums and benefits on the life cycle. Most LTC coverages today are fully funded, and the age at the inception of the contract is closely related to the expected duration in good health which is a key factor for the premium-benefit ratio. Calendar time reflects cohort effects in the risk distributions. In a fully funded setting, the age distribution of the portfolio does not influence the rating structure. Hence a run-off decision should not have a specific effect on the rating structure.

- The age of the policyholder (sum of the seniority and of the age at the inception of the contract) is a key factor for the premium-benefit ratio of an unfunded coverage without cross-subsidies between age classes, such as an annual renewable term life insurance. Calendar time is related to cohort effects. As death rates increase by 8-9 percent each year in middle age, similar values would be expected in the regression for seniority and for the age at the inception of the contract, if the coverage was rated in this context. Figures of eight or nine percent are well above the 1.5 and 0.5% values obtained for the death benefit coverage. Hence strong cross-subsidies are expected between age classes for the life insurance component.

- Unfunded financing with cross-subsidies between age classes is usual in health insurance. In an unfunded setting, calendar time is important because the premium-benefit ratio reflects the current age distribution. Calendar time is the only variable that matters if the coverage is rated according to a "community rating" principle.

Table 2
Rating structure of the three coverages

	Death benefit	Health	Long-term care
R^2	0.585	0.927	0.861
Intercept	-6.013 (**)	-2.498 (**)	-4.394 (**)
Female gender	0.015 (**)	0.001 (**)	0.006 (**)
Age at inception	0.015 (**)	$<10^{-4}$ (**)	0.018 (**)
Seniority	0.005 (**)	$<10^{-4}$ (**)	0.008 (**)
Cancellation year			
1993	-1.037 (**)	-0.482 (**)	-0.793 (**)
1994	-1.014 (**)	-0.385 (**)	-0.778 (**)
1995	-1.007 (**)	-0.290 (**)	-0.769 (**)
1996	-1.008 (**)	-0.225 (**)	-0.772 (**)
1997	-1.016 (**)	-0.138 (**)	-0.768 (**)
1998	-0.932 (**)	-0.091 (**)	-0.723 (**)
1999	-0.807 (**)	-0.051 (**)	-0.616 (**)
2000	-0.676 (**)	-0.052 (**)	-0.511 (**)
2001	-0.524 (**)	-0.052 (**)	-0.370 (**)
2002	-0.389 (**)	-0.052 (**)	-0.237 (**)
2003	-0.252 (**)	-0.054 (**)	-0.114 (**)
2004	-0.105 (**)	-0.058 (**)	-0.015 (**)
2005	-0.088 (**)	-0.058 (**)	-0.012 (**)
2006	-0.059 (**)	-0.026 (**)	0.005

Note: Logarithm of the premium-benefit ratio, explained by a linear model including gender, the age at the inception of the contract, the seniority of the policyholder and binary variables related to the year with available premiums (2007 for a working contract, or the cancellation year). The sample contains 150,123 contracts working or cancelled between 1993 and 2006. Reference levels: gender=male; cancellation year=2007 or working contract. Significance level for the nullity test: Two stars indicates a p-value less than 0.01; One star corresponds to a p-value comprised between 0.01 and 0.05.

There is almost no gender effect in the rating structure of the death benefit coverage. The calendar effect is very important, which suggests that the coverage is not fully funded.[10] A Lee-Carter (1992) analysis shows that we have mortality improvements, as is the case in the whole of the Spanish population. We would observe a decreasing calendar effect in a fully funded setting. Instead, the calendar effect remains almost constant between 1993 and 1997, and then increases sharply once the portfolio is set in a run-off position (the ratio is multiplied by $\exp(1.016) = 2.76$ from 1997 to 2007). This result can be partly explained by the presence of term life insurance. Between 1993 and 1997, policyholders could age with very low increases in the death benefit premium, due to the continuous arrival of new, young policyholders. This was stopped by the run-off decision, which reflects the insurance company's lack of commitment with respect to portfolio renewal. On the other hand, there is partial funding because of the surrender value, and there is a significant effect of the age at the inception of the contract, as for fully funded schemes.

As indicated in Table 2, the premium-benefit ratio of the health coverage depends only on calendar time. When controlling for this variable, there is no residual effect of gender, age at inception or seniority in the portfolio. The mutual company follows a "community rating" strategy for health coverage. The calendar effect is different here from the effect observed in the death benefit coverage. The rating level significantly increases before the run-off decision, but afterwards it becomes more stable. As can be seen in Table 5, there is a cohort risk improvement which explains the stability of the calendar effect in a context of continuous aging.

Results for long-term care are close to those obtained for death benefit insurance, with the same conjecture of partial funding due to the importance of calendar effects after the run-off decision. However, these calendar effects could be partly explained in a funded setting by an increase in future LTC risks due to mortality improvements.

The following sections analyze the cross-subsidies between age classes for each of the coverages, which determine the magnitude of front-loading and are of interest in the analysis of lapse behavior. Health risk is analyzed in greater depth in order to estimate reclassification risk through the history of disability spells.

[10] As we do not know the reserves, the results that follow for the type of funding are merely conjectural.

4.2 Death benefit coverage

For each coverage, we derived a benefit-premium ratio in the portfolio for the period 1993 to 2006, from the rating structure estimated in Table 2.[11] Due to the low level of premiums and the individual nature of the policies, there is a high ratio between management costs and premiums. Therefore, the loading factor (equal to one, minus the benefit-premium ratio) is high. Death benefit insurance is more heavily loaded than the two other coverages because there is a surrender value, which was not available to us and could not be taken into account in our benefit-premium derivations. The benefit-premium ratios given in the following tables are expressed with respect to an undisclosed average. There is no gender effect in the rating structure, and women are younger than men on average (44 years vs. 56 years, partly due to the absence of women before the 70s). Hence women strongly subsidize men, as regards the death benefit coverage. The benefit-premium ratios with respect to the global average are equal to 1.27 and 0.21 respectively for men and women.

A comparison of the age effect on risks and premiums clearly indicates that young policyholders subsidize older policyholder's death benefits. Table 3 presents the benefit-premium ratios compared to the global average, when policyholders are grouped by decades.

Table 3

Benefit-premium ratios compared to the global average

Age class (years)	<30	[30,40[	[40,50[	[50,60[	[60,70[	[70,80[	≥ 80
Benefit-premium ratio (w.r.t. average)	10%	11%	24%	39%	91%	222%	598%

4.3 Health coverage

Disability risk includes both a frequency and a duration component, and a comprehensive statistical approach assesses the two aspects separately. We restrict ourselves here to a semiparametric analysis of the prevalence in the disability state, which should suffice to estimate the risk borne by the insurance company. Then we assess the predictive ability of the disability history on different components of health risks.

[11] Benefits and premiums are updated each year according to an inflation index. We extended the last available premium to the preceding years on the basis of this index.

First, we present some global statistics regarding the frequency of disability spells, disability prevalence and the benefit-premium ratio expressed with respect to the global average. These statistics are presented according to the age classes used in Table 3.

Table 4
Global statistics on health risks

Age class (years)	Frequency of disability spells	Disability prevalence	Benefit-premium ratio (w.r.t. average)
< 30	0.105	1.09%	45.1%
[30, 40[	0.140	1.75%	57.8%
[40, 50[	0.185	3.00%	70.7%
[50, 60[	0.213	4.80%	93.3%
[60, 70[	0.214	7.18%	132.1%
[70, 80[	0.207	9.36%	175.6%
≥ 80	0.166	10.87%	224.0%
whole population	0.180	4.65%	100%

The first two global results suggest an average duration of three months for a disability spell. The annual frequency of disability spells increases with age up to 60-70 years, and then decreases. However, disability prevalence increases during the whole life cycle, and the duration of disability spells increases steadily with age. Young policyholders subsidize older ones, as shown in the last column. This result is typical for individual health insurance.[12] Disability prevalence is multiplied by ten when comparing the oldest policyholders with the youngest ones. By contrast, the benefit-premium ratio is multiplied by five for the same age classes although this ratio depends solely on calendar time. The link between the benefit level and disability prevalence, explained as a function of age, is important for understanding this discrepancy. A more detailed analysis shows that this link is strong and positive for young policyholders (i.e. young policyholders entitled to large benefits have a higher disability prevalence), but weakens with age. A positive link is expected because policyholders might forget to use their coverage in the case of disability, especially if the benefit level is low. As for the

[12] Pauly and Herring (1999) estimate an elasticity of premiums with respect to changes in expected expenses due to age in a sample of American contracts. They report results that range from 0.2 to 0.45.

gender effect, the benefit-premium ratios with respect to the global average are 0.88 for women and 1.04 for men, which is a more equitable result than in the case of death benefits.

We estimate disability prevalence with a logit link in a generalized linear model. This approach needs to be justified, as the dependent variable belongs to the interval $[0, 1]$, but is not binary. Suppose that the transition intensity from good health to disability is equal to λ, and that the duration of a disability spell is a random variable D. The stationary disability prevalence r is equal to

$$r = \frac{E(D)}{E(D) + \frac{1}{\lambda}} = \frac{1}{1 + \frac{1}{\lambda E(D)}} = \frac{1}{1 + \exp(-s)} = \Phi(s), \ s = \log(\lambda) + \log E(D). \quad (1)$$

The limit rate given in equation (1) is a logistic function of a score s which can be expressed as a linear form of regression components. This is the usual specification for $\log(\lambda)$, and is also the case for $\log E(D)$ if the distribution family for the duration is of the "accelerated life models" type. A random variable ranging in $[0, 1]$ has a variance inferior to that of a binary variable with the same expectation. However, we use the link between variance and expectation of binary variables in our generalized linear model estimation. The estimators are more accurate than what the regression would indicate, but they are consistent if the expectation is well specified. Table 5 summarizes the regression results.

The statistical units are contract-years, and the dependent variable is the disability duration. Table 5 can be interpreted as follows. The disability rates are much less than one, and the corresponding scores s are negative. As we have $\Phi(s) \simeq \exp(s)$ at the neighborhood of $-\infty$, a small absolute variation in the score s is related to a similar relative variation in the corresponding expectation $\Phi(s)$. For instance, women are roughly 20% riskier than men *ceteris paribus*, contrary to the gender averages. This discrepancy is not surprising given that, on average, women are much younger than men. The cohort effect is clearly decreasing. Disability prevalence risk increases with age, with a slight downturn at the end of the life cycle when the other variables are controlled. The link between the benefit level and disability prevalence is positive but decreases with age, as discussed earlier.

Table 5
Disability prevalence estimated with a logit link

Regression component	Estimation
Intercept	-4.2251 (**)
Age class (years)	reference level: $\geq$ 80 years
$<$ 30	-3.7571 (**)
[30, 40[	-2.8465 (**)
[40, 50[	-2.0730 (**)
[50, 60[	-0.7867 (**)
[60, 70[	-0.0207
[70, 80[	0.2401 (**)
Gender	reference level: male
Female	0.2015 (**)
Calendar year	reference level: 2006
1993	1.0654 (**)
1994	1.1152 (**)
1995	1.0713 (**)
1996	0.9929 (**)
1997	0.9042 (**)
1998	0.7859 (**)
1999	0.7221 (**)
2000	0.6818 (**)
2001	0.6435 (**)
2002	0.5691 (**)
2003	0.5047 (**)
2004	0.3645 (**)
2005	0.1652 (**)
log(daily benefits)*age class	
$<$ 30 years	0.5051 (**)
[30, 40[	0.3945 (**)
[40, 50[	0.3208 (**)
[50, 60[	0.2034 (**)
[60, 70[	0.1660 (**)
[70, 80[	0.1999 (**)
$\geq$ 80 years	0.3059 (**)

Note: disability duration for a contract-year (1,182,662 individuals), explained by a logit model where covariates are age classes, gender, calendar year and the logarithm of daily benefits crossed with age classes. The p-values are represented as in Table 2.

The history of disability spells is a key variable for understanding health status. The disability history allows symmetric learning, but the insurance company is committed not to use this information in its rating structure. We now assess the predictive ability of the disability history on different components of health risks. Experience rating models use distribution mixing[13], but such an approach is not straightforward for transition models between different health states. Rather, we integrate an offset variable based on a "bonus-malus" coefficient in the regression of Table 5.[14] We retain a coefficient of the type

$$BM = \frac{a + d}{a + \widehat{E}(D)}, \tag{2}$$

where d is the cumulated past duration in the disability state, and where $\widehat{E}(D)$ is its estimated expectation from the regression detailed in Table 5. This bonus-malus coefficient is similar to that usually found in frequency risk models (with number of claims instead of durations), and the weight given to the individual health history decreases with the parameter a. The logarithm of the bonus-malus coefficient is the offset variable, and the estimated prevalence is almost proportional to BM for the average individual, because $\Phi(s) \simeq \exp(s)$. If a policyholder has a better health history than her peers (i.e. if $d < \widehat{E}(D)$ with the covariates used in Table 5), the offset variable is negative, and there is a health "bonus". We estimate a with likelihood maximization.[15] The likelihood increases dramatically when the offset variable is included in the regression of Table 5. We obtain $\widehat{a} = 0.0984$. To illustrate, we assess the influence of a supplementary year on an average individual. The corresponding values are equal to

$$d = \widehat{E}(D) = 0.24;\ s = \Phi^{-1}(0.0465) = -3.02.$$

The average disability prevalence is equal to 4.65%, and the average disability duration corresponds roughly to a five year risk exposure. Suppose that there

[13] See Pinquet (2000) for a survey.

[14] In a regression, the parameter related to an offset variable is set equal to one.

[15] The log-likelihood is derived from the logit model on binary variables. Applied on prevalence rates, it can be seen as a dissimilarity index of the Kullback-Leibler type between observed and estimated values.

is a disability spell with an average duration (i.e. three months) during the following year. The bonus-malus coefficient varies from one to

$$(0.0984 + 0.24 + 0.25)/(0.0984 + 0.24 + 0.0465) = 1.53.$$

The score s increases by 0.42 if we leave the regressors unchanged, and the corresponding estimated disability prevalence increases by 49%. If the policyholder remains healthy during the whole year, the estimated disability risk decreases by 11%. These two values should be corrected by the consequences of the policyholder aging and by calendar effects.

Disability history also helps predict LTC and death risks. The following results reinforce the reasons to lapse the contract because of a good health history. We estimate proportional hazards models on the age at entry into an LTC spell and on the age at death, with gender and $\log(BM)$ as covariates. The estimated elasticities of the death rate and of the entry rate into LTC with respect to the bonus-malus disability coefficient are equal to 0.477 and 0.622 respectively. Using the average disability history, a supplementary year with a three month disability spell entails a 22% increase in risk of death and a 30% increase in risk of entry into LTC. These results should also be corrected for aging and cohort effects.

4.4 LTC coverage

The LTC risk in this portfolio has been analyzed in an earlier article (Guillén and Pinquet, 2008). Table 6 presents statistics for prevalence and benefit-premium ratios. Cross-subsidies between age classes are very strong (i.e. young policyholders subsidize older ones), in a similar way to the death benefit coverage. Buying LTC coverage in one's 30s (which is the average age at purchase: See Table 1) is very uncommon, and can only be explained by the bundled nature of the product.

Table 6
Statistics for LTC coverage

Age class (years)	LTC prevalence	Benefit-premium ratio (w.r.t. average)
< 30	0.10%	4.1%
[30, 40[	0.22%	19.4%
[40, 50[	0.31%	33.5%
[50, 60[	0.64%	51.5%
[60,70[	1.43%	99.8%
[70, 80[	3.22%	230.0%
≥ 80	7.83%	576.2%
whole population	1.09%	100.0%

5 Empirical results on lapses

In the introduction and Section 2 of this paper, we stressed the important redistribution effect of lapses, which motivates an empirical analysis of lapsation behavior in the portfolio. Let us detail first the different variables of interest in this analysis. We assessed the magnitude of front-loading in Section 4, and front-loading deters lapses if surrender benefits are low. Lapse rates increase with the magnitude of reclassification risk, which is represented by the health "bonus-malus" coefficient in the following derivations. The evolution of the age structure of the portfolio influences lapse behavior in an unfunded setting. Thus, an increase in lapse rate is expected after the run-off decision taken in 1997. Lapse decisions may also be related to wealth (either liquidity constraints or loss of interest in the contract on the part of affluent policyholders). Finally, the policyholder's understanding of the contract at the date of purchase may influence lapse behavior, as we discuss below.

The average health bonus-malus coefficient derived previously is presented in Table 7 for lapsing policyholders, who are grouped by decade. Unsurprisingly, policyholders who cancel their contract have good health histories compared those of to their peers, and this difference is found to increase with age. Reclassification risk increases with seniority, and seniority also increases with age. Hence lapsation for reasons of risk is less likely for young policyholders. Policyholders who lapse below the age of 40 are only a few percentage points less risky than their peers with respect to the three risks covered in the bundle (see

Table 7 and the elasticities estimated in Section 4.3). Hence reclassification risk is unlikely to be a major reason for lapsation among young policyholders.

Table 7
Health history for lapsing policyholders.

Age class (years)	<30	[30,40[	[40,50[	[50,60[	[60,70[	[70,80[	≥80
Percentage of Cancelled contracts	24.7%	21.8%	19.5%	15.1%	13.8%	4.2%	0.9%
Average Health BM coefficient	0.968	0.903	0.869	0.867	0.847	0.704	0.562

Note: There are 66,451 cancelled contracts. Health Bonus-Malus coefficient: $BM = \frac{0.0984+d}{0.0984+\widehat{E}(D)}$. The average bonus-malus coefficient for lapsing policyholders is equal to 0.888.

A detailed analysis of lapsation behavior is provided in Table 8. We estimate a proportional hazards model for the age of the policyholder, where events are lapses. The covariates include gender, the logarithm of the health bonus-malus coefficient, and the year of risk exposure.

The elasticity of the lapse rate with respect to the health "bonus-malus" coefficient is negative as expected and very significant (the limit level in a test for nullity is negligible). The gender effect is small, and the calendar effects have the same shape as the gross surrender rates given in Table 1. Note that the company did not offer an alternative contract to policyholders when the portfolio was closed to new business in 1997, but it is unclear as to whether this decision was publicized or not. However, the behavior of the policyholders in the ten years following this run-off decision shows that many of them did not behave optimally. Indeed, the run-off decision is detrimental to all the policyholders that remained in the portfolio, because of induced supplementary aging and partial funding. Middle-aged policyholders are most negatively affected by this decision. They subsidized the portfolio when they were young, and if they decide to lapse they are repaid only a fraction of their claim. If they choose to stay however, they will not benefit from subsidies provided by new policyholders. It is difficult to see why young policyholders should stay in the portfolio after the run-off decision. They would obviously find better conditions with a new contract. Yet, Table 1 shows that, after 1997, the portfolio is not pulled towards the "death spiral" that could have been expected after a continuous departure of the youngest policyholders.[16] This did not occur, and we believe that the portfolio reaches an equilibrium thanks to a mistaken perception of the situation by the youngest policyholders. As discussed previously, they are probably unaware of the run-off decision, nor of their position in the age distribution of the portfolio.

Table 8

Proportional hazards model for lapsation

Covariate	Parameter estimate	Hazard ratio
$\log(BM)$	-0.091	
Gender	Reference level: male	
Female	-0.020 (*)	0.980
Year of risk exposure	Reference level: 1993	
1994	0.173 (**)	1.189
1995	0.343 (**)	1.409
1996	0.327 (**)	1.386
1997	0.618 (**)	1.855
1998	0.707 (**)	2.028
1999	0.679 (**)	1.972
2000	0.667 (**)	1.949
2001	0.656 (**)	1.926
2002	0.691 (**)	1.995
2003	0.490 (**)	1.632
2004	-0.373 (**)	0.689
2005	-0.386 (**)	0.680
2006	-0.326 (**)	0.722

Note: proportional hazards estimation of cancellation risk. Number of observations: 1,163,645 duration-event indicator pairs. The basic observation is a contract-year, because the health "bonus-malus" coefficient is updated every year. Durations are left truncated by the age at the beginning of the year. The covariates include gender, the logarithm of the health bonus-malus coefficient, and year of risk exposure. The p-values are represented as in Table 2.

[16] See Pauly, Mitchell and Zeng (2007) for an analysis of "death spirals" in health insurance.

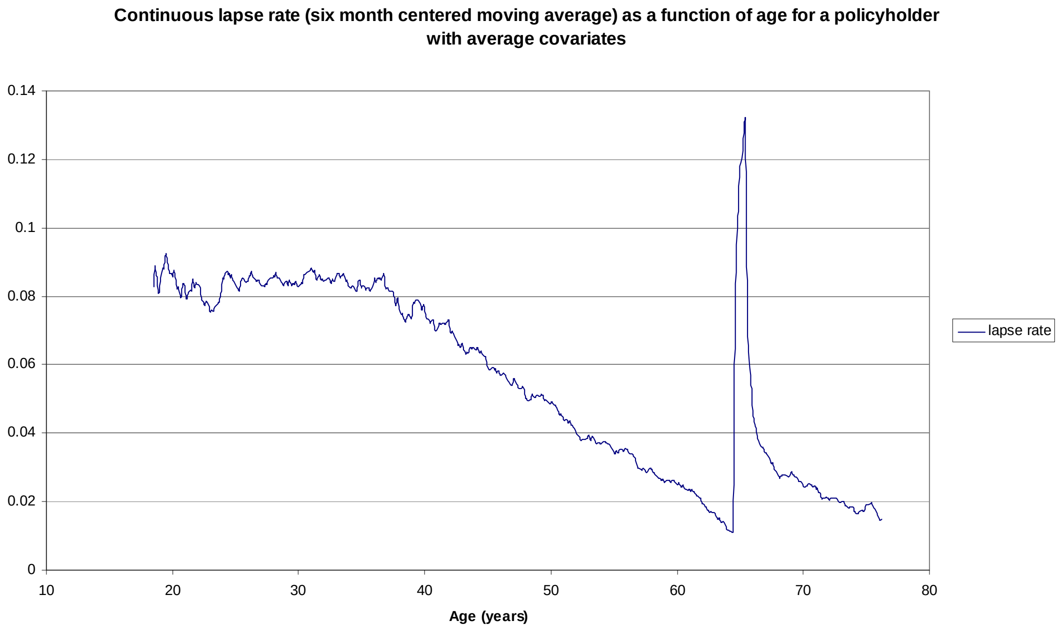

Figure 1:

Finally, the baseline hazard function (i.e. the continuous lapse rate derived for a policyholder with the average covariates, as a function of age) is shown in Figure 1.

Lapse rates decrease after 35 years, but there is a local peak at 65 years. Benefits are not modified in any way at the date of retirement, but the policyholders may have perceived this wrongly, which might account for this peak.[17]

The decreasing link between age and lapse rate has several explanations. First, lapse motivations induced by liquidity constraints decreases with wealth, and wealth increases with age. Second, the front-loaded nature of the contract

[17] Another reason is that retired people are not any more subject to the risk of income loss caused by a disability spell.

leads to the same outcome. The losses incurred by lapsation increase with seniority in the portfolio, and younger policyholders have a lower seniority. Hence they are less deterred by losses if they want to lapse for whatever reason. Due to the nature of the guarantees, young policyholders are more likely to switch to another contract. A sum of 1,600 Euros (the average death benefit) is scarcely enough to cover the costs of a funeral, and young policyholders may cancel their policy as they redefine their insurance needs. For instance, they might switch to term life insurance if their need for family protection increases. This would be more rational than combining a whole life insurance, a health insurance and an LTC coverage, as the latter is usually bought by much older people. Let us recall that once this portfolio was closed to new business, the mutual company no longer sold this type of bundled product. Hence a misunderstanding of the product or changes in insurance motives can also explain the high lapse rate of young policyholders. From Figure 1, these arguments supporting higher lapse rates among younger policyholders outweigh the reclassification risk argument which works at the opposite, as discussed after Table 7.

These results show that a knowledge of the structure of the portfolio or of the guarantees influences lapse rates. Long-term insurance purchasing is also particularly sensitive to an awareness of the environment, as discussed by Zhou-Richter, Browne and Gründl (2010). On the basis of a survey performed in Germany, they show that demand for LTC insurance is low due to an underestimation of the risks and costs related to LTC. The more aware adult children are of the risks and costs, the more likely LTC insurance is to be purchased, either by the children themselves on behalf of their parents or by the parents under the influence of their children.[18]

Other causes of lapsation should also be borne in mind. Table 1 shows evidence of aging in the population that lapses after the run-off decision, with a sharp increase between 2003 and 2004. This aging goes in tandem with a decrease in the lapse rate. Elderly people may lapse for reasons of health or wealth. A good state of health is more and more informative as age increases, and this is expressed in the health bonus-malus coefficient (the denominator increases with time, as it is related to risk exposure). The results in Table 7 suggest that lapse motivations for reasons of health increase with age. However, wealth arguments may also explain this trend since some policyholders may be affluent enough to lose interest in the low benefits associated with the policy.

[18] In their sample, the willingness to buy LTC insurance increases from 1.8% to 30.7% after an information update on the risks and costs of LTC.

Most Spanish households own their home, and older homeowners are quite likely to have repaid their home mortgages. Thus, they are in a position to convert home equity into annuities that pay for medical expenses. Davidoff (2009) shows that if consumers can liquidate home equity in the event of severe illness, LTC insurance as well as annuities become less attractive.

6 Conclusions

Based on this case study, we believe that insufficient knowledge of insurance products can cause lapsation which is detrimental to policyholders if triggered by information available at the date of purchase. Enhancing knowledge of the environment in terms of risk and of insurance solutions would be welfare improving. It would also increase insurance demand, as argued by Zhou-Richter, Browne and Gründl (2010).

In this case study, LTC coverage was not fully funded. Otherwise the rating structure would not depend on the age structure of the portfolio and would not have been influenced by the run-off decision. Nowadays, LTC coverages are fully funded and this is beneficial to policyholders in open portfolios.

A run-off decision is detrimental to policyholders in an unfunded setting if cross-subsidies are substantial, especially if the insurance company follows a "community rating" policy. This issue concerns health insurance first and foremost. In such a context, regulating authorities should prevent the splitting of portfolios, so that the insurance company is committed to keeping its portfolio open.

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