ECOLE POLYTECHNIQUE

CENTRE NATIONAL DE LA RECHERCHE SCIENTIFIQUE

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April 2009

Cahier n° 2009-13

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Résumé: Cet article présente une étude de cas sur un portefeuille de contrats d'assurance de long terme. Nous décrivons les niveaux de risque, la structure de tarification et les subventions croisées induites sur le portefeuille pour un contrat joint de trois garanties santé, dépendance et vie entière. Nous mettons en évidence le risque de reclassification à partir de l'historique des épisodes de maladie. Nous analysons le comportement de résiliation et tentons d'expliquer la dynamique observée sur le portefeuille. Enfin, nous tirons des conclusions sur la conception des contrats et les différences entre assurance privée et assurance publique.

Abstract: This paper presents a case study of a portfolio of individual long-term insurance contracts. We describe the risk levels, the rating structure and the implied cross-subsidies on the portfolio of a bundle of three coverages related to health, life and long-term care. We show evidence of reclassification risk through the history of disability spells. We also analyze the lapse behavior and try to give a rationale for the observed dynamics of the portfolio. Lastly, we draw conclusions regarding the design of such contracts and the difference between private and public insurance.

Classification JEL: C01, C23, D12.

Mots clés : Engagement, risque de reclassification, assurance à long terme.

Key Words : Commitment, reclassification risk, long-term insurance.

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1 Introduction

This paper presents a case study of a portfolio of individual long-term insurance contracts, which were sold by a Spanish mutual insurance company (the portfolio has been set in a run-off position for ten years). The contract is a bundle of three coverages related to health, long-term care (referred to later as LTC) and life.[1] As is usual with long-term contracts, there is a one-sided commitment in terms of loyalty. The policyholders can leave the company, but the latter cannot cancel the contract. The policyholder is then insured against reclassification risk, because experience rating is also forbidden. Lack of long-term insurance (whether public or private) against health or life risks may entail important welfare losses (Diamond, 1992). However the insurer is not committed to a long-term premium scheme, and the average premium level follows the average loss trend. If the ratio premium-benefit 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 important calendar effects. Due to mortality improvements, the companies benefit from cohort effects as regards death benefit insurance (whether term life or whole life insurance). However, as a result of aging, long-term care frequency risk increases with calendar time. A natural hedge of an insurance company against uncertainty in a Knightian sense is not to commit to a long-term premium scheme. In our study, the premium-benefit ratio for each coverage depends on calendar time, and also on age and seniority (except for the health component of the bundle). The age effect in the rating structure is much lower than the link between age and short-term risk for each component of the contract, and gender is not taken into account. These characteristics entail strong cross-subsidies between genders, generations and also between the periods of a contract. Besides, there is a surrender value for the death benefit component of the bundle, but none for the LTC coverage.

Issues related to commitment, cross-subsidies between periods and lapse behavior in long-term insurance contracts have already been addressed extensively by economic literature. Cross-subsidies between the periods of a

[1] In the first 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 the 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.

contract are termed as "lowballing" or "highballing", depending on whether the first periods are subsidized by the following ones, or the contrary.[2] The contracts analyzed in the paper are of the "highballing" type. Young policyholders, although they pay less than the older ones, strongly subsidize them as will be shown in Section 3. "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). Such information could be obtained in our data from the history of disability spells, but experience rating is forbidden in the long-term contract analyzed in the paper. Fluet, Schlesinger and Fei (2009) discuss multiperiod contracts with a "lowballing" structure and an opting out opportunity, which is interest when the motivation to insure varies with time.[3] As argued by Pauly, Kunreuther and Hirth (1995), risks that evolve in unpredictable ways (such as health and life risks) are more 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 are Hendel and Lizzeri (2003), with an empirical study on term life insurance linked with a model with symmetric learning, one-sided commitment and buyer heterogeneity in the cost of front-loading. The model predicts that low risks may prefer a multiperiod contract with front-loading (i.e. "highballing"). This result is confirmed empirically by the average premiums observed in the USA on 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 any, is often very low as compared to a retrospective actuarial value.[4] From a statistical

[2]Alternative terminology is "back-loading" and "front-loading".

[3]The authors propose a term life insurance contract designed to cope with changes in bequest motives. The contract is sold at a subsidized price in the first period, then policyholders discovering a low bequest motive in the second period have an option to buy back the contract at an unfair price. In their framework, the contract achieves a higher welfare level than the naive strategy of delaying insurance purchase until one's bequest needs are known.

[4]An exception is the whole life insurance coverage. Besides, these contracts are assignable, which allows the creation of secondary markets. The associated arrangements are termed "life settlements" and "viatical settlements". See Doherty, Singer (2002), and

study of long-term care American contracts, Brown and Finkelstein (2007, 2009) derive an average loading factor equal to 0.18 if the policy is held until death, whereas the loading factor rises to 0.51 if lapses are accounted for. To our knowledge, empirical results on the causes of lapsation in long-term insurance contracts are not available. If the main reason for cancelling the contract is a liquidity constraint, we would have unpleasant redistribution effects because poor policyholders would subsidize richer ones.

The article is organized as follows. Section 2 describes the insurance contract, and the portfolio analyzed in the empirical study. We also present the economic framework, and in particular the evolution of public and private health insurance in Spain. The three risks covered by the insurance bundle are assessed in Section 3. Our study focuses on health risk, because life risk is well known and LTC risk for this portfolio is analyzed in a previous paper (Guillén and Pinquet, 2008). Another reason is that the history of disability spells is key in the learning on the policyholder's health status, and hence may influence lapsation behavior. We show evidence of reclassification risk through the history of disability spells. We also analyze the rating structure of the company and the implied cross-subsidies. We find that the mutual company strictly follows a "community rating" strategy for the health coverage, and that young policyholders strongly subsidize the older policyholders. We provide empirical evidence on the lapse behavior in Section 4, and try to give a rationale for the results. We find that policyholders who cancel their contract have good health histories compared to their peers, and that the lapse rate decreases with age, with a local peak at 65 years. We argue that lapse behavior of young policyholders and also of elderly policyholders at retirement is partly due to a misunderstanding of the contract. We also comment the fact that the portfolio avoids the "death spiral" that could have been expected after the run-off decision from a continuous departure of the youngest policyholders. Section 5 comments the design of long-term contracts and draws conclusions on the difference between private and public long-term insurance.

Daily, Hendel and Lizzeri (2008) for the pros and cons of secondary life insurance markets.

2 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 of 2007. Due to modifications in the computing environment in 1992, the contracts that were closed before that date were suppressed from the data base. Although data are available from 1975, we restrict our analysis from 1993 to 2006 to contracts that are either working or cancelled after 1992. The contract is a bundle of three coverages, namely a death benefit insurance, a health coverage and a long-term care component. There is no age constraint for the benefits, and all coverages extend into the whole lifespan regardless of the employment status of the insured, who can either be employed, in retirement or in another 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 company, but the latter cannot cancel the contract. Hence attrition is due only to cancellation at the initiative of the policyholder, or to death. 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 LTC spells related benefits 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 levels for premiums and benefits indicate that the product plays a role of additional coverage. They also fit the financial means of policyholders who are young at the inception of the contract (see Table 1). Benefits are indexed on an inflation rate. The ratio premium-benefit varies over time without precommitment (see Section 3 for results). Benefits and premiums cannot be modified on an individual basis, neither for risk reasons (experience rating) nor for modifications of the coverage level. The insurance package can only be cancelled completely, and only the death benefit component has a surrender value.

The health coverage works when the policyholder temporally requires medical treatment and cannot perform daily activities in the meantime. The state of disability is assessed by doctors appointed by the company on the basis of standard medical and physical tests. There is no connection between the compensation decision made by the company and the decisions made 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, who are not able to perform daily life activities without the assistance of another person. Conditions to become eligible are very 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, due to the traditional practices of the company, blindness or loss of two arms or legs are sufficient conditions to grant compensation. Finally, death benefit does not have constraints on 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 build up a fund to compensate women facing an untimely death of their husband. Later, death benefit coverage was extended in order to meet the needs of a policyholder during the life cycle, as seen by the company agents. People underwriting the product were not called "insured" but "affiliated". This product became popular among workers, who were exclusively men at the time. Women were not affiliated until the 70s, because in Spain they did not participate in the job market before that decade. From that time on, the Spanish government instituted the so-called "Development Plans" (Planes de Desarrollo) which fostered social and economic change and allowed women to start entering the job market. A real public health and welfare system was also created in the late 70s. Initially, it 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 the paper had roughly 60000 affiliates in 1960. That number increased to 170,000 in 1975 and to 250,000 in 1984. The product originally was distributed in Catalonia, which was the company's main area of influence. During the 80s, the product was also sold in other Spanish regions as the company expanded. However, the company stopped selling the product from 1997, and from that time on, the coverages were sold separately.

Let us give some basic statistics on the portfolio dynamics. In Table 1 we present on average and for each calendar year: the age of policyholders (x), the 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 percentage of the portfolio size).

Table 1
Descriptive statistics for the portfolio

Average age of policyholders, 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)

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

We observe a steady aging of the portfolio, but the rate and causes of this aging vary with time. If the variables of the table were defined in continuous time, the time derivative of the average age would be equal to

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

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 decreases afterwards. The most striking result 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.

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

3.1 The rating structure

The rating structure of the three components of the insurance package is described in Table 2, based on the analysis of the premium-benefit ratio. Premiums and benefits recorded in the data base are the last values available, and relate either to the cancellation year or to 2007 for working contracts. Indicators of the collection year for premiums and benefits are included in the regression, in order to assess the calendar effects in the rating structure. The main driving force of the rating level is calendar time, which means that the rating policy is almost of the "community rating" type.

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	0	0.018
Seniority	0.005	0	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 inception, the seniority of the policyholder and binary variables related to the year with available premiums (2007 for a working contract, or the cancellation year). Reference levels: gender=male; cancellation year=2007 or working contract.

There is almost no gender effect in the rating structure of the death benefit coverage. For a given policyholder, a supplementary year entails a 0.5% increase in the premium-benefit ratio (due to the seniority variable), whereas a supplementary year of age at inception entails a 1.5% increase of the ratio. The calendar effect is very important. The overall rating level remains almost constant between 1993 and 1997, and then increases sharply once the portfolio is set in a run-off position. The obvious explanation is the continuous aging of the portfolio induced by the run-off decision. Policyholders could age between 1993 and 1997 with very low increases in the death benefit premium, due to the continuous arrival of new and young policyholders. This was stopped by the run-off decision, which reflects the insurance company's lack of commitment with respect to the portfolio renewal.

The health benefits in the ratio analyzed in Table 2 are derived on a monthly basis in order to maintain the same periodicity as that of the premiums. As indicated in that table, the premium-benefit ratio of the health coverage depends only on calendar time. Controlling for this variable, there is no residual effect of age, gender and seniority in the portfolio. The mutual company follows a "community rating" strategy for health coverage. The calendar effect is here different from the effect observed in the death benefit coverage. The rating level significantly increases before the run-off decision, but afterwards it is more stable.

Results for long-term care are close to those obtained for death benefit insurance.

3.2 Death benefit coverage

Let us first describe the death benefit insurance component of the bundle. Death rates have the usual properties (i.e. the risk level for women is twice lower than for men for a given age and cohort, and the annual rate of increase is equal to 8-9% for a given cohort and gender, at least in middle age). Calendar effects are negative during the whole period (from 1993 to 2006), reflecting a continuous mortality improvement. For instance, a proportional hazards model on death risk where gender, birth date and entry date in the portfolio are included as regressors provides a 1.7% discount for each supplementary birth year. If age, gender and birth cohort are controlled, the seniority in the portfolio does not have any effect on death risk. A positive link between seniority and death risk was expected, because low risks among peers are more expected to cancel the contract.

For each coverage, we derived a benefit-premium ratio on the portfolio during the 1993-2006 period, from the rating structure estimated in Table 2.[5] 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 hence was not 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 almost no gender effect in the rating structure, and women are younger than men on average (44 years vs. 56 years, partly due to women's absence before the 70s). Hence women strongly subsidize men, as regards the death benefit coverage. The benefit-premium ratios are equal to 0.21 and 1.27 with respect to the global average.

A comparison of the age effect on risk and premium obviously suggests that young policyholders subsidize the 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
Relative ratio	0.10	0.11	0.24	0.39	0.91	2.22	5.98

The strong cross-subsidies between the age classes and the existence of a surrender value make the death benefit component actually function as a whole life insurance. The ratio increases more than expected at the end of the life cycle because women are absent from the portfolio beyond the eighty years threshold. Again, this is a consequence of their inclusion only since the 70s.

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

3.3 Health coverage

Disability risk includes both a frequency and a duration component, and a comprehensive statistical approach assesses the two aspects separately. We shall restrict to a semiparametric analysis of the prevalence in the disability state, which is enough to estimate the risk borne by the insurance company. First, we present in Table 4 some global statistics on the frequency of disability spells, disability prevalence and the benefit-premium ratio expressed with respect to the global average. These statistics are given 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 until 60-70 years, and then decreases. However disability prevalence increases during the whole life cycle. Hence 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.[6] Disability prevalence is multiplied by ten if the oldest policyholders are compared to the youngest ones. On the other hand, the benefit-premium ratio is multiplied by five for the same age classes. This result may seem surprising at first glance, as the rating level of the health coverage only depends on calendar time, as shown in Table 2. Due to the run-off position

[6] Pauly and Herring (1999) estimate an elasticity of premiums with respect to changes in expected expense due to age on a sample of American contracts. They find results ranging from 0.2 to 0.45.

of the portfolio, oldest policyholders are related to more recent periods than younger ones. The average rating level increases with calendar time, which partly explains the discrepancy between the two last columns of the table. However the link between the benefits level and disability prevalence, explained as a function of age, is even more important for the understanding of 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. This positive link between the benefit level and the disability prevalence can be explained by adverse selection or moral hazard arguments. In addition, policyholders may forget to use their coverage in case of disability, especially if the benefit level is low. This last result will be used in the estimation of the disability prevalence which follows. Concerning the 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. Lastly, let us mention that the benefit-premium ratio decreases steadily with calendar time.

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

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

The limit rate given in equation (1) is a logistic function of a term 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. 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.

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

The statistical units are contract-years, and the dependent variable is the disability prevalence for each given year. The logistic function Φ is such that $\Phi'/\Phi = 1 - \Phi$. An absolute variation of the score s which is small with respect to one is related to a similar relative variation in the corresponding expectation $\Phi(s)$ if the latter is close to zero. To put it differently, we have $\Phi(s) \simeq \exp(s)$ at the neighborhood of $-\infty$, which leads to the preceding conclusion. This is the case on average in this regression. For instance, women are roughly 20% riskier than men *ceteris paribus*, contrary to the averages on gender. This discrepancy is not surprising as, on the average, women are much younger than men. The calendar effect is clearly decreasing. Disability prevalence risk increases with age, with a slight downturn at the end of the life cycle when other variables are controlled. The link between the benefit level and disability prevalence is positive but decreases with age, as previously mentioned.

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[7], 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 preceding regression. We retain a coefficient of the type

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

where d is the cumulated past duration in the disability state, and where $\widehat{E}(D)$ is its estimation from the regression detailed in Table 5. This bonus-malus coefficient is similar to the one 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. We estimate a with likelihood maximization. The likelihood of the binary model is seen here as an adjustment coefficient between observed and estimated prevalence values,[8] and it increases dramatically when the bonus-malus coefficient is included in the regression. The logarithm of the bonus-malus coefficient is the offset

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

[8] The log-likelihood can be expressed from dissimilarity indexes of the Kullback-Leibler type.

variable, and the estimated prevalence is almost proportional to BM for the average individual, because then $\Phi(s) \simeq \exp(s)$. We obtain $\widehat{a} = 0.0984$. To illustrate, let us assess the influence of a supplementary year for 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 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. We estimate proportional hazards models on the age at entry into LTC 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 respectively equal to 0.477 and 0.622. 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 be also corrected by aging and calendar effects.

3.4 LTC coverage

The LTC risk has been analyzed in a preceding article (Guillén and Pinquet, 2008). We analyzed the entry rate into LTC from 1975 to 2005, but unfortunately we were unaware that contracts cancelled before 1992 were suppressed from the data base. However, results on the entry rate after 1992 remain largely the same even after correcting for this oversight. The entry rate into LTC increases by 12 percent with each supplementary year of age during most of the life cycle. A Lee-Carter analysis shows that there is no significant calendar effect on the entry rate between 1993 and 2006[9]. As for the duration of LTC spells, we obtain the following results if we control for gender, age and date at the beginning of the LTC spell. *Ceteris paribus,* a supplementary year of age increases the mortality rate by 4.9%, and a supplementary calendar year is associated with a 1.8% increase. The mortality rate of women in the LTC state is 15.7% lower than that of men, if the two other variables are controlled for. The calendar effect is much lower than that given in our preceding article, and the joint age and calendar effects for an individual along a Lexis line (i.e. the line in a plane age-calendar time followed by a given cohort) are lower than the increase in the entry rate. Hence the prevalence of the LTC state increases with age, a well known result which means that aging will entail a growth in the frequency of LTC spells. Table 6 presents statistics on prevalence and on benefit-premium ratios.

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%

Cross-subsidies between age classes are very strong, similar to the death benefit coverage. Buying of a LTC coverage in one's 30s (the average age at purchase: See Table 1) is very uncommon, and can only be explained by the bundled nature of the product. Besides, the benefit-premium ratio increases with calendar time during the first years, then remains stable. Its average value lies between that of the death benefit and of the health coverage.

[9] See Guillén and Pinquet (2008) for more details on the model. The Lee-Carter (1992) specification is applied to a transition intensity from good health to an irreversible LTC state instead of an observed rate, as is usually the case.

4 Empirical evidence on lapses

General results on lapses were presented in Table 1. The estimation of disability prevalence shown in Table 5 also allows for an appraisal of the individual history of the policyholder, and of its influence on her current status in the portfolio. The health bonus-malus coefficient derived previously is presented on average in the Table 7, depending on the current status of the policyholder.

Table 7
Health history and current status of the policyholder

Current status of the policyholder	$\overline{BM}$	Frequency
Cancelled policy	0.943	51.5%
Good health, or in temporary disability	0.883	38%
In an LTC state	1.971	0.7%
Dead after an LTC spell	1.588	1.5%
Dead, without a previous LTC spell	1.866	8.3%

Note: Bonus Malus coefficient: $BM = \frac{0.0984+d}{0.0984+\widehat{E}(D)}$

Not surprisingly, policyholders who cancel their contract have good health histories compared to their peers, but the difference is small. A more detailed analysis of lapsation behavior is provided in Table 8. We estimate a proportional hazards model on the age of the policyholder, where events are lapses. The covariates include gender, the logarithm of the health bonus-malus coefficient, and 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 rough surrender rates given in Table 1. The calendar effect increases before 1997, which might have incited the company to offer a new contract. This effect remains at a high level until 2003, and then it drops suddenly between 2003 and 2004. It is worth noticing that the company did not offer an alternative contract to policyholders when the portfolio was closed to new business in 1997. It is not clear whether the run-off decision was publicized or not. The surrender rate increased from 1996 to 1997, whereas the rating structure varied as in the preceding year (see Tables 1 and 2), so some policyholders might have heard of the run-off decision. However, their behavior in the ten years that followed proves that many of them did not behave optimally, as we discuss later.

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

Lastly, 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. The decreasing link between age and lapse rate comes as no surprise, since older policyholders in the portfolio benefit more from cross-subsidies. Benefits are not modified in any way at the date of retirement, but this fact may be wrongly perceived by the policyholders, possibly explaining this peak.

Let us discuss the possible causes of lapsation. 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

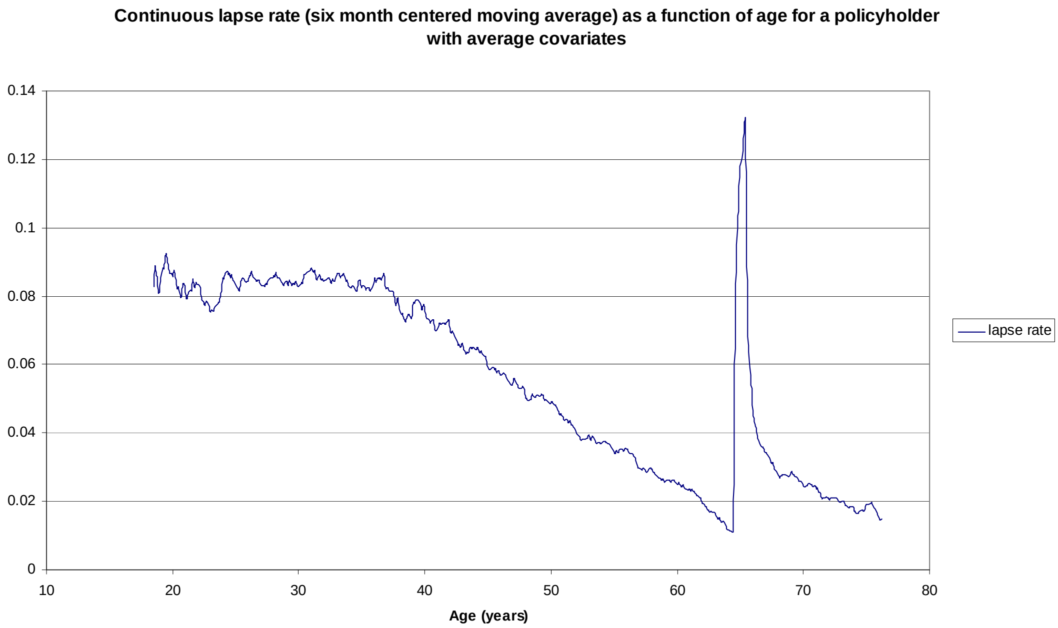
Continuous lapse rate (six month centered moving average) as a function of age for a policyholder with average covariates
0.14
0.12
0.1
0.08
0.06
0.04
0.02
0
10
20
30
40
50
60
70
80
Age (years)
lapse rate

Figure 1:

in the lapse rate. Elderly people may lapse for health or wealth reasons. A good health state is more and more informative as age increases. We mentioned an average health bonus-malus coefficient of 0.943 in Table 7 for the cancelled contracts. Depending on the cancellation year, we observe a drop of the bonus-malus coefficient between 2003 and 2004 (from 0.88 to 0.79), which suggests that health reasons may motivate lapsation among the elderly. 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. Most Spanish households own their home, and older homeowners have most likely repaid their loans. Thus, they can 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.

The low value of death benefits can also be unattractive for younger people. A sum of 1,600 Euros (the average death benefit) is hardly enough to finance a funeral, and young policyholders may cancel their policy because they redefine their insurance needs. For instance, they could focus on term life insurance if their need of 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 did not sell this type of bundled product any more. The last argument for cancellation was the liquidity constraint, but it should have had a marginal effect since the premiums were low. As a conclusion, either a misunderstanding of the product or changes in insurance motives can explain the high lapse rate of young policyholders.

The run-off decision taken in 1997 is detrimental to all the policyholders remaining in the portfolio, as supplementary aging entails an increase in all premiums. 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 will be repaid for a tiny part of their claim as the contract is mostly unfunded, except for a part of the death benefit coverage. If they decide to stay, they will not be part of the oldest population in the portfolio and will not benefit later from cross-subsidies. It is difficult to find a reason why young policyholders should stay in the portfolio. They would obviously find better conditions with a new contract. Table 1 shows however that, after the run-off decision, the portfolio is not pulled towards the "death spiral" that could have been expected after a continuous departure of youngest policyholders.[10] This has not happened, and we believe that the portfolio reaches an equilibrium due to a mistaken perception of their situation by the youngest policyholders. As already mentioned, they are probably not aware of the run-off decision.

5 Conclusions

In the conclusion we compare the analyzed insurance product to a public insurance coverage in an attempt to point to its advantages and drawbacks. A basic argument in support of private insurance coverage is that it better matches the heterogeneity of preferences. Lump-sum coverages, as in life insurance, are appropriate for the satisfaction of these various needs. This argument also supports the European policy with regard to long-term care insurance, where the benefit level is defined *ex ante*.

Private coverages have on average a higher loading factor than public ones, and the size of the insured group is paramount in this respect. Diamond (1992) mentions a 40 percent loading factor for groups with a size of fewer than 5, and a 5.5 percent loading factor for groups of 10,000 and more. In our portfolio of individual policies, the loading factor is not surprisingly closer to the first figure quoted by Diamond.

Long-term insurance lacks readability. This characteristic applies to all coverages, whether public or private, but the freedom of choice left to consumers in the purchase of private long-term insurance leads to decisions that are often seen afterwards as suboptimal. For instance, the high lapse rate of young policyholders suggests that the purchase decision is often regretted a few years later. On the other hand, many policyholders do not leave the company after the run-off decision probably because they are not aware of it. In addition, many of them probably ignored that the insurance company could take such a decision when they purchased the coverage. The surrender loss, which represents the discounted accumulated premiums in excess of the corresponding liabilities and of the surrender value, generates a risk which is probably underestimated by the consumers in their purchase decision of long-term insurance. Regulating authorities should enforce a better readability of long-term insurance contracts, which would help the consumers to make better decisions and would also improve the competitive environment.

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

References

[1] Brown, J., and A. Finkelstein, 2007, Why is the Market for Long-Term Care Insurance so Small?, *Journal of Public Economics* 91 (10): 1967–1991.

[2] Brown, J., and A. Finkelstein, 2009, The Private Market for Long-Term Care Insurance in the United States: A Review of the Evidence, *The Journal of Risk and Insurance* 76 (1): 5–29.

[3] Daily, G., I. Hendel, and A. Lizzeri, 2008, Does the Secondary Life Insurance Market Threaten Dynamic Insurance?, *American Economic Review, Papers & Proceedings* 98: 151-156.

[4] Davidoff, T., 2009, Housing, Wealth, and Annuities,*The Journal of Risk and Insurance* 76 (1): 31–52.

[5] Diamond P., 1992, Organizing the Health Insurance Market, *Econometrica* 40: 1233-1254.

[6] Dionne, G., and N. Doherty, 1994, Adverse Selection, Commitment, and Renegotiation; Extension to and Evidence from Insurance Markets, *Journal of Political Economy* 102: 209-235.

[7] Doherty, N., and H. Singer, 2002, The Benefits of a Secondary Market for Life Insurance Policies, Wharton School working paper.

[8] Fluet, C., H. Schlesinger, and W. Fei, 2009, Uncertain Bequest Needs and Long-Term Insurance Contracts, working paper.

[9] Guillén, M., and J. Pinquet, 2008, Long-Term Care: Risk Description of a Spanish Portfolio and Economic Analysis of the Timing of Insurance Purchase, *The Geneva Papers on Risk and Insurance - Issues and Practice* 33: 659-672.

[10] Hendel, I., and A. Lizzeri, 2003, The Role of Commitment in Dynamic Contracts: Evidence from Life Insurance, *The Quarterly Journal of Economics* 118(1): 299-327.

[11] Kunreuther, H., and M. Pauly, 1985, Market Equilibrium with Private Knowledge: An Insurance Example. *Journal of Public Economics* 26: 269-288.

[12] Lee, R.D., and L. Carter, 1992, Modelling and Forecasting the Time Series of US Mortality, *Journal of the American Statistical Association* 87: 659-671.

[13] Pauly, M., and B. Herring, 1999, Pooling Health Insurance Risks. The AEI Press.

[14] Pauly, M., H. Kunreuther, and R. Hirth, 1995, Guaranteed Renewability in Insurance, *Journal of Risk and Uncertainty* 10(2): 143-156.

[15] Pauly, M., O. Mitchell, and P. Zeng, 2007, Death Spiral or Euthanasia? The Demise of Generous Group Health Insurance Coverage, *Inquiry* 44(4): 412-427.

[16] Pinquet, J., 2000, Experience rating through heterogeneous models, in: G. Dionne, Ed., *Handbook of Insurance* 459-500.

[17] Webb, D.C., 2009, Asymmetric Information, Long-Term Care Insurance, and Annuities: The Case for Bundled Contracts, *The Journal of Risk and Insurance* 76 (1): 53–85.