Sunk Costs and Screening: Two-Part Tariffs in Life Insurance*

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Abstract

There are large, upfront, fixed costs to writing a life insurance policy. Both agent commission and direct underwriting costs (e.g., fees for physicals and blood tests) are fully paid a few years into contracts that can last 10-30 years. Because of these upfront costs, insurers can actually lose money on policies when the consumer lapses early into the contract, even if no death benefit is ever paid out. Thus, to properly price contracts, insurers must estimate lapse risks. However, consumers will often have private knowledge of their lapse likelihood, leading to adverse selection. We develop a model of insurance pricing under heterogeneous lapse rates with asymmetric information about lapse likelihood within the context of an optional two-part tariff as a screening device for future policyholder behavior. We then test for consumer self-selection using detailed, policy-level data on life insurance backdating (a common practice that resembles a two-part tariff). We are able to identify, through a control function approach, the information about lapse risk a consumer reveals when they choose to backdate. Our contribution to the literature is twofold: we are the first to consider life insurance lapsing as a form of adverse selection; we also explore, both theoretically and empirically, the role of optional two-part tariffs as a screening mechanism using life insurance backdating as our primary example. We find that consumers who are less likely to lapse self-select into the two-part tariff pricing structure and also document consumer behavior consistent with sunk cost fallacy.

JEL Codes: **D80, G11, G22**

Keywords: **Asymmetric Information, Adverse Selection, Screening, Sunk Cost Fallacy, Insurance Markets, Life Insurance, Lapsing**

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1. Introduction

There are large, upfront, fixed costs to writing a life insurance policy. Both agent commission and direct underwriting costs (e.g., fees for physicals and blood tests) are fully paid a few years into contracts that can last 10-30 years. Because of these upfront costs, insurers can actually lose money on policies when the consumer lapses early into the contract, even if no death benefit is ever paid out. Thus, to properly price contracts, insurers must estimate lapse risks. However, because consumers often have more knowledge about their lapse likelihood than the insurer, asymmetric information arises and room for a screening mechanism exists. In this paper, we develop a model of life insurance pricing under heterogeneous lapse behavior with asymmetric information about lapse likelihood. We establish the existence of a separating equilibrium under a menu of contracts containing an optional two-part tariff. We then show the consumer's choice serves as a screening device for private information on lapse likelihood. Using detailed, policy-level data on life insurance backdating as an example of our proposed optional two-part tariff, we empirically test our model's prediction of consumer self-selection.[1] We use a control function approach to separately identify selection effects from potential sunk cost fallacy. We find that consumers who choose to take part in the two-part tariff by backdating their policies are less likely to lapse, due to both self-selection and sunk costs.

Lapse rates in life insurance are substantial. Between 1991 and 2010, $29.7 trillion of new individual life insurance coverage was issued in the United States. During this same time period, $24 trillion of coverage lapsed. 85% of term life insurance policies never pay a death benefit. More surprisingly, 74% of term life policies sold to people at age 65 fail to pay a death benefit (Gottlieb and Smetters, 2014).[2] In a given year, for every policy that ends in death or term maturation, 36 policies lapse due to nonpayment of premiums (Purushotham, 2006). We focus our analysis on

[1] Backdating occurs when the contract's start date is earlier than its application date. Most often, policies are backdated to align financial documents or set a specific premium payment date. However a smaller number (6% of all policies in our sample) are backdated to save age, which is described in detail later. Throughout this paper we are only considering policies backdated to save age.

[2] Whole life insurance, which doesn't expire, has similar overall lapse rates and lower per-year lapse rates.

term life insurance and, though the vast majority of literature on insurance lapsing is focused on whole life insurance, much of the theory applies to term insurance as well. The literature on lapse behavior is well-developed and can be condensed into: preference shocks, income shocks, policy replacement, and non-expected utility explanations.

Preference shocks refer to any number of situations where the private value of the life insurance contract has changed (Fang and Kung, 2010; Liebenberg et al., 2012; Fei et al., 2015).[3] Examples include: divorce, death of a spouse or child, children becoming self-sufficient, increase in spousal income, etc. Income shocks refer to consumers experiencing a negative shock to income and thereby having insufficient funds to pay premiums. The effect of an income shock on lapse rates is stronger in whole life insurance due to the higher premiums and presence of a surrender value. This is referred to in the literature as the emergency fund hypothesis (Linton, 1932; Outreville, 1990; Kuo et al., 2003). The aptly named policy replacement hypothesis refers to consumers who lapse on one policy because they find what they believe to be a better one (Outreville, 1990; Carson and Forster, 2000). The interest rate hypothesis is a specific case of the policy replacement hypothesis where the driver of better available policies is a change in expectations of future interest rates (Schott, 1971; Pesando, 1974; Kuo et al., 2003).[4] The final category of research on lapse rates focuses on non-expected utility models of consumer behavior and how these various behavioral assumptions can influence the decision to lapse on a policy (Shefrin, 2002; Mulholland and Finke, 2014; Gottlieb and Smetters, 2014).[5] Our theory requries that consumers potentially have some private knowledge about their lapse proclivity. Of these four main determinants, consumers most likely have private knowledge about income shocks and preference shocks.

Though asymmetric information on lapse risk initially appears different than the canonical study of asymmetric information on loss probability, the same intuition applies. In our case, we observe that life insurance incurs a front-loading of underwriting-related costs both through direct underwriting costs and agent commission. It is expected that the costs of underwriting will be recovered by the insurer over the long life of the contract. However, when a consumer lapses on

[3]This reason for lapsing generally only covers negative shocks to the insurance value of the policy. Positive shocks to preferences are typically subsumed by the policy replacement category.

[4]Because insurance premiums are collected, and invested, long before benefits are paid out, expectations about interest rates play an important part in the determination of premium rates.

[5]For an excellent, detailed analysis of these hypotheses the interested reader is referred to Eling and Kochanski (2013).

a policy early into the contract (e.g., in the first several years), the insurer is unable to recoup the entire underwriting cost. Thus, consumers with a higher likelihood of lapsing have a higher (expected) average cost per year because the fixed cost of underwriting is spread over a smaller time span. If consumers have private knowledge about this cost, the same adverse selection issues arise.

Much has been written on possible asymmetric information issues in life insurance, though the extant literature focuses solely on loss as opposed to lapse. Cawley and Philipson (1999) show the existence of a negative correlation between risk and coverage in life insurance that runs counter to the testable implications set forth by Chiappori et al. (2006). Cutler et al. (2008) explore reasons for this apparent "advantageous selection" citing a negative correlation for risk aversion and risky behaviors. Gottlieb (2012) suggests it is unlikely there is an adverse selection problem in life insurance, as it is priced close to actuarially fair levels. However, the possible presence of simultaneous advantageous selection, which produces opposite observable correlation as adverse selection, clouds this analysis (De Meza and Webb, 2001). Under advantageous selection, individuals who have lower risk also tend to be more risk averse. This produces a situation where less risky individuals are more likely to purchase insurance due to their higher risk aversion even though they have a lower probability of experiencing a loss.[6]

The main difference in lapse risk versus loss risk is that the common methods of addressing asymmetric information do not readily apply: deductibles and copays will not cause consumers to self-select into their proper lapse-risk group. Instead we show insurers can offer a menu of contracts where the variety lies in combinations of an initial fee and recurring premiums, i.e., a menu of two-part tariffs. We show that consumers who choose to pay the higher up-front fee and lower recurring premiums are signaling their intention to persist and not to lapse early in the contract period.

Two-part tariffs, originally examined by Gabor (1955) and Bowman (1957), are central to the literature concerning price discrimination. The simplest form of second-degree price discrimination, two-part tariffs are commonly used across numerous industries. The first formal analysis of this form of price discrimination was Oi (1971) which examines the optimal pricing structure for amusement parks (Disneyland in particular); Schmalense (1981) expands on this work. Blackstone (1975) looks into potential two-part tariffs for Electrofax copying machines. Mussa and Rosen (1978) was the

[6]For a broad discussion of adverse and advantageous selection, see Einav and Finkelstein (2011).

firsd to examine two-part tariffs as a screening mechanism. Schmalense (2015) looks at the "razor-and-blades" pricing strategy.[7] The extreme version of a two-part tariff is "buffet pricing," where the per-period price is set to zero and the entire payment is subsumed by the tariff (Nahata et al., 1999). Prior work examined on why life insurance contracts tend to be actuarially front-loaded, but the extant literature has failed to consider that optional front-loading acts as a screening device (Hendel and Lizzeri, 2003; Hofmann and Browne, 2013).[8]

A version of an optional two-part tariff pricing method exists in the life insurance industry, called backdating. A life insurance contract is considered backdated when the insurance contract bears a start date that is prior to the application date.[9] The consumer chooses to pay for coverage for the time prior to their application and, because the consumer is still alive at the time of application, no death benefit will be paid for this prior coverage. Rather, consumers do this to "save age." Because life insurance policies feature level premiums that are based largely on the age of the applicant at the beginning of the contract, backdating to save age allows the consumer to be underwritten at a lower age, which lowers the per-period premium paid throughout the life of the contract.[10] Throughout this paper we are only considering policies backdated to save age. The choice to backdate is, in essence, the choice to pay an initial upfront payment to have lower per-period payments, which looks a lot like an optional two-part tariff. The extant literature on backdating is limited. This practice is quite common. In our sample, 19% of policies legally able to be backdated are.[11] Carson (1994) shows that the net present value of backdating can be positive after relatively short periods of time depending on the length of the backdate period, the discount rate, and age. Carson and Ostaszewski (2004) further shows that the actuarial present value of backdating is generally positive. Carson et al. (2012) examine the incentives and welfare economics of life insurance backdating.

Prior work on life insurance backdating has failed to address if and why decreased lapse rates occur. The obvious avenue through which greater policy persistence might occur is via a selection

[7] For a more in-depth discussion of the extant literature on two-part tariffs, the interested reader is referred to Stole (2007); for a discussion of the theory the reader is referred to Tirole (1988).

[8] There is a somewhat similar vein of work on reclassification risk in health insurance (Herring and Pauly, 2006; Pashchenko and Porapakkarm, 2015; Handel et al., 2015).

[9] This process is regulated by states with typical maximum backdating of one year.

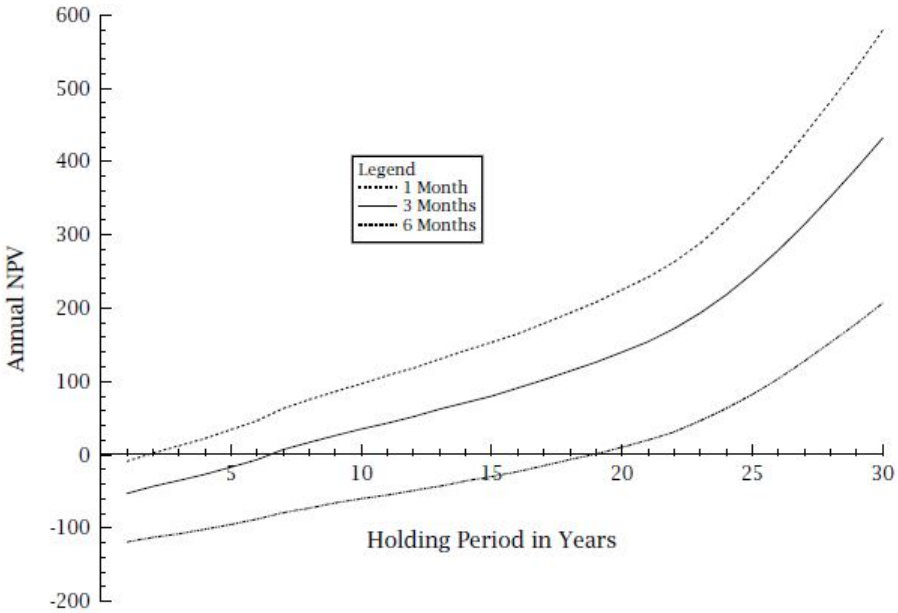
[10] We note also that other forms of two-part tariffs could be offered by insurers (in addition to backdating), such as the option to pay for underwriting /policy issue expenses separately from mortality charges, for example.

[11] The laws vary from state to state, with the most common restriction being a maximum backdate length of 6 months.

mechanism – those consumers with asymmetric knowledge of their low lapse risk are more likely to pay the initial tariff because the future stream of decreased premiums is longer and, therefore, more valuable to them. Behavioral economics offers a different mechanism: sunk cost fallacy (Arkes and Blumer, 1985).[12] Sunk cost fallacy would lead to lower lapse rates because consumers may have an aversion to "wasting" the high upfront tariff of backdating. If, for instance, consumers are randomly assigned to the tariff pricing structure, we would not expect those who pay the tariff to exhibit lower lapse rates, unless those consumers are exhibiting sunk cost fallacy (and we are properly controlling for the differences in per-period premiums). Through our empirical structure we are able to separately identify the self-selection and behavioral effects of backdating on lapsing behavior. Our work builds on the growing field evidence of behavior consistent with sunk cost fallacy (Ho et al., 2017).

Figure 1

NPV of Backdated Policy Relative to Normal Policy



Note: Annual net present value (NPV) for a 35 year old male for a $250,000, 30 year term policy. Source: Carson and Ostaszewski (2004)

Because of the discrete nature of life insurance underwriting, specifically the use of integer ages, the relative distance from a consumer's application date to their birthday will change the size of

[12] We acknowledge the existence of potential rational explanations for consumers exhibiting what appears to be sunk cost fallacy (McAfee et al., 2010). We are only documenting the existence of the behavior, not necessarily providing evidence of irrational behavior.

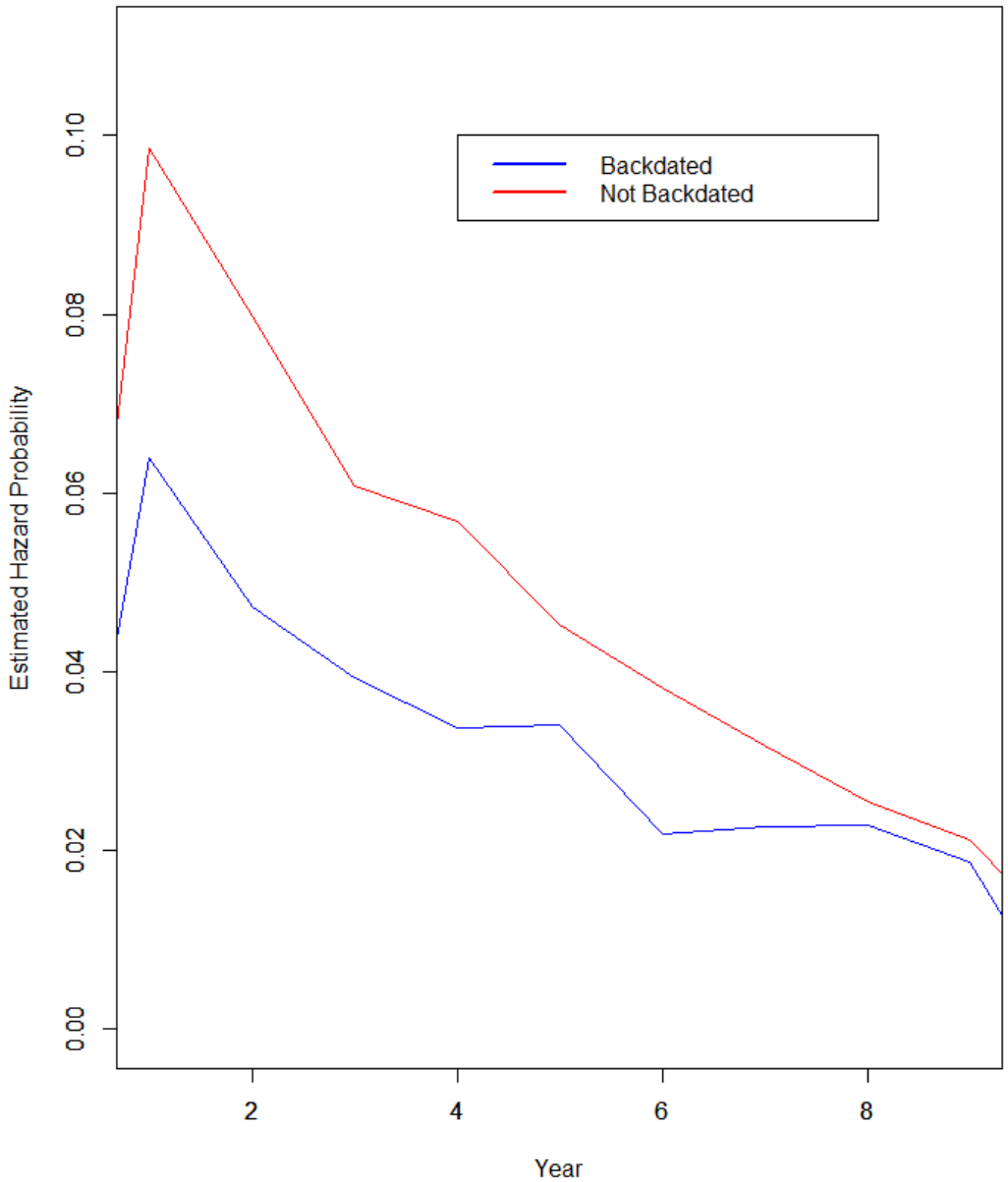
the initial tariff while having no effect on the size of the premium reduction. Drawn from Carson and Ostaszewski (2004), Figure 1 shows how the actuarial value to a 35-year-old male consumer of backdating a policy evolves over time compared to a traditionally purchased policy for different lengths of backdating. Consumers who apply for life insurance far away from their birthdays must pay a larger upfront cost to get the same reduction in premium. Thus, it takes longer for the premium reduction to outweigh the initial tariff. We exploit this randomness in initial tariff size to separately identify the selection and sunk cost effects of backdating on lapse proclivity. We find each avenue to be a significant contributor to a consumer's decreased propensity to lapse.

Our paper is similar to Gottlieb and Smetters (2014). We examine the dilemma, to the insurance company, that the consumer will lapse on their contract too early – Gottlieb and Smetters (2014) examine the potential they will lapse too late. Both of these dilemmas are due to level premiums. Ours arises because most costs incurred in acquiring contracts are heavily frontloaded which means, due to level premiums, the contract must be in force for at least a few years for the insurance company to profit from the contract. Gottlieb and Smetters's (2014) arises because, due to death risk increasing with age, level premiums are actuarially frontloaded. This means that in the later parts of contracts the annual expected payout is higher than the annual premium. At this point, and thereafter, the insurance company is better off if the consumer lapses on the contract. Furthermore, Gottlieb and Smetters's (2014) stated reason for consumers selecting into contracts they are likely to eventually lapse on, consumer over confidence in retention rates, is not mutually exclusive from our examination. We only require that different types of consumers have different expectations of their own lapse rates. We do not require these expectations be unbiased. Our results of consumer selection likely do not have an effect on Gottlieb and Smetters's (2014) results because, as Figure 2 shows, as the effect seems to fade away as time passes.

Our contribution to the literature is twofold: we are the first to consider life insurance lapsing as a form of adverse selection; we also explore, both theoretically and empirically, the role of optional two-part tariffs as a screening mechanism using life insurance backdating as our primary example. The findings shed light on why insurers continue to engage in backdating when alternative pricing mechanisms exist – backdating is used by insurers as a screening device to deal with hidden knowledge of lapse intentions that engender a form of adverse selection. As part of our results, we also show evidence of consumer behavior consistent with sunk cost fallacy.

Figure 2

Difference in Hazard Probabilty over Time



Note: Difference in hazard probability over time for backdated vs. non-backdated policies.

The rest of this paper proceeds as follows: the next section describes a simple model of life insurance pricing under heterogeneous lapse behavior with perfect asymmetric information about lapse likelihood and then examines the role of optional two-part tariffs as a screening device. The

third section describes our data and empirical methods, and the fourth section presents our results. The final section offers conclusions.

2. Theory

We begin with a simple expression of how insurers can use optional two-part tariffs to screen a consumer's lapse risk.[13] Let there be two periods. Consider a term insurance contract offered by a competitive life insurer lasting both periods with level premiums. Let there be two types of consumers: those who lapse after the first period and those who continue into the second. λ is the share of consumers who lapse on their policies after the first period and $(1-\lambda)$ is the share who do not lapse. Both types of consumers have an equal chance ρ_1 of loss L in the first period and ρ_2 of loss L in the second period. Define $\rho_2 > \rho_1$ to exemplify that death risk increases with age. Consumers are indistinguishable to the insurer who must incur a fixed underwriting cost $F > 0$. The profit Π for a representative insurer, separated by period, offering full coverage at level premium P is

$$\Pi = \Pi_1 + \Pi_2$$

$$\Pi_1 = P - \rho_1 L - F$$

$$\Pi_2 = (1-\lambda)P - (1-\lambda)\rho_2 L$$

Where Π_t is the profit in period t.

With a competitive insurer, we use the zero profit condition and solve for equilibrium price

$$P^* = \frac{F + \rho_1 L + (1-\lambda)\rho_2 L}{(2-\lambda)}.$$

The first derivative of premium with respect to lapse rate is then

$$\frac{\partial P^*}{\partial \lambda} = \frac{F + (\rho_1 - \rho_2)L}{(2-\lambda)^2}$$

[13] In this section we are ignoring any irrational behavior, specifically sunk cost fallacy.

The sign of this derivative, which determines whether or not lapsing is costly to the insurer, is determined by the relative size of the underwriting cost F and the difference in expected costs between periods $(\rho_1 - \rho_2)L$. In situations such as shorter term life insurance written for younger consumers, it is likely that $F > (\rho_1 - \rho_2)L$ and thus $\frac{\partial P^*}{\partial l} > 0$ meaning lapsing is costly to insurers. In other situations, such as longer term insurance written to the elderly, $(\rho_1 - \rho_2)L$ may be quite large and lapsing may actually reduce the costs of the insurer. In either case, the optimal price for the two types differs and room for a screening mechanism exists.

If insurers instead offer an optional two-part tariff, the equilibrium changes. A pooling equilibrium is impossible under the standard Rothschild and Stiglitz (1976) argument so we instead search for a separating equilibrium. This equilibrium is defined by two separate premium offers (P_1, P_2) where $P_1 > P_2$ and an initial tariff T that a consumer must pay to access P_2.[14] This equilibrium space is defined by a solution to the following:

$$P_1 = \rho_1 L + F \tag{1}$$

$$2P_2 + T = (\rho_1 + \rho_2)L + F \tag{2}$$

$$2P_2 + T \leq 2P_1 \tag{3}$$

$$P_2 + T \geq P_1 \tag{4}$$

Equations (1) and (2) represent the zero profit conditions; Equations (3) and (4) represent the incentive compatibility constraints. A space of solutions always exists, with an intuitive example being to set the tariff T as slightly higher than the underwriting cost $F + \epsilon$ and P_2 as slightly lower than the expected, first period, loss $\rho_1 L - \epsilon$, where ϵ is small.[15]

[14] Because our model is only two periods, a two-part tariff for those consumers who lapse after one period is indistinguishable from a single premium.

[15] Unlike Rothschild and Stiglitz (1976), a solution always exists in our model.

3. Data and Methods

3.1. Data

The data for the analysis come from a medium-sized and geographically diverse mutual life insurer in the U.S. and spans contracts written in the years 2005 through 2013.[16] The full sample includes data for 97,522 term life insurance contracts. We focus exclusively on term insurance between 10-30 years.[17] Because death claims are so rarely paid in the first nine years of term policies, the potential for consumers holding their policies "too long" appears remote. That is, with reference to our model, it is likely that $F > (\rho_1 - \rho_2)L$, thus making improved persistency strictly a benefit for the insurer and separating our analysis from Gottlieb and Smetters (2014).[18] This allows us to view higher likelihood of lapse strictly as an adverse form of selection and adhere more to the predicted selection results of our model. To avoid the complication of classifying deaths in our hazard model, we remove the 727 policies that end in death.[19] Many states have laws restricting how long a policy can be backdated. Of the 10 states our insurer operates in – 5 restrict backdating to six months, 1 restricts to 3 months, and 4 have no laws. We remove from the analysis every application that is greater than 6 months from the applicant's birthday.

Table 1: Summary Stats

Variable	Full Sample		Backdate		Don't Backdate	
	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Backdate	0.19	0.39	1	-	0	-
Backdate Opp	86.72	47.74	44.04	23.44	96.70	46.46
Face Amount($000s)	223.79	299.47	312.19	431.72	203.13	254.62
Issue Age	41.13	11.11	43.29	10.91	40.63	11.10
Annual Premium	584.22	1,036.33	874.21	1,764.73	516.44	756.95
Lapse	0.30	0.46	0.21	0.41	0.32	0.47
Days in Force	1,508.17	1,020.75	1,736.21	1,012.505	1,454.87	1,015.34
N	24,661		4,672		19,989	

Note: This table shows summary statistics separated by consumers who did and did not backdate.

[16] We use these years to simplify the issue of a term policy expiring vs. a consumer lapsing.

[17] This focus eliminates the yearly renewable term contracts. As these policies are spot contracts and do not have level premiums, a two-part tariff has no meaning. 10-year terms are the smallest length term policies in our data after 1-year renewable term. This also eliminates universal and whole life contracts.

[18] In our full sample of 97,522 contracts, we observe 727 deaths; in our selected sample, we observe 494.

[19] Inclusion of policies that end in death and classifying them as non-lapsed policies does not affect our results. Similarly, classifying the policies as lapsing at death does not affect our results.

Life insurance policies are often backdated for reasons other than saving age (such as to align financial documents). To address this, we define a policy as backdated only if the effective date is within one week prior to the consumer's birthday. Our identification strategy relies on randomness in applying for life insurance with regard to a consumer's birthday. However, there is evidence that consumers do alter their behavior during the short time frame around their birthday (Dai et al., 2014). To address this we drop those consumers who apply within one month (before or after) of their birthday.[20] After these modifications, and the removal of missing data, our sample consists of 46,199 active or lapsed policies. Summary statistics for these policies can be found in Table 1. Because they have to pay a lower tariff for the same reduction in premiums, consumers who apply for life insurance closer to their birthdays are more likely to backdate their policies. Additionally, consumers with backdated policies both hold their policies longer and have lower lapse rates. Interestingly, consumers who backdate tend to have larger policies.

3.2. Base Specification

To first check our proposed optional two-part tariff selection mechanism, we examine if backdated policies exhibit lower lapse rates. Our initial specification for the hazard of lapsing is defined as follows:[21]

$$\begin{aligned}
\lambda(Lapse_{it}|X) &= \lambda_0(t)exp(X_i\beta')u_i \qquad (5)\\
X_i &= (\mathbf{Backdate_i}, IssueAge_i, Male_i, FaceAmt_i, AnnPrem_i,\\
&\qquad UWClass_i, TermLength_i, IssueYear_i, State_i)\\
u_i &\sim exp(N(0,\sigma^2)).
\end{aligned}$$

We fit Equation (5) using the partial likelihood method of Cox (1972) $Lapse_{it}$ is equal to 1 in time t if the policy has lapsed at any time equal or prior to t and 0 otherwise. For the independent

[20] If we include these consumers our results are slightly strengthened.

[21] Contrary to most hazard models, our data are not truly panel and thus we do not represent them as such. Each policy represents one row in our data and we derive the hazard specifications from a combination of our time-invariant covariates: length of time policy was/is in force and a binary indicator for if the policy lapsed. Since our data's time frame for policy issuance is smaller than the minimum term length used in our data (10 years) we circumvent the need to account for issues with a policy expiring without lapse or policy-holder death.

variables in the model, $Backdate_i$ is 1 if the consumer was found to be saving age and 0 otherwise. $IssueAge_i$ is the age of the insured at time of application. $FaceAmt_i$ is the face value of the life insurance policy. $AnnPrem_i$ is the annual premium for the policy. $UWClass_i$ is a series of dummy variables for the underwriting class of the insured (with four dummy variables for five different categories), with $UWClass1_i$ being the healthiest. $Male_i$ is a binary variable equal to 1 if the insured is male, and 0 otherwise. $Year_i$ is a series of dummy variables for the year of policy issuance. $TermLength_i$ is a series of dummy variables for length of term (10, 15, or 20 years). $State_i$ is a series of dummy variables for state of issuance. Because we are going to use a control function in the next step, we additionally specify the structure of the error term u_i as being log-normally distributed with mean 0 and variance σ^2.

3.3. Control Function Approach

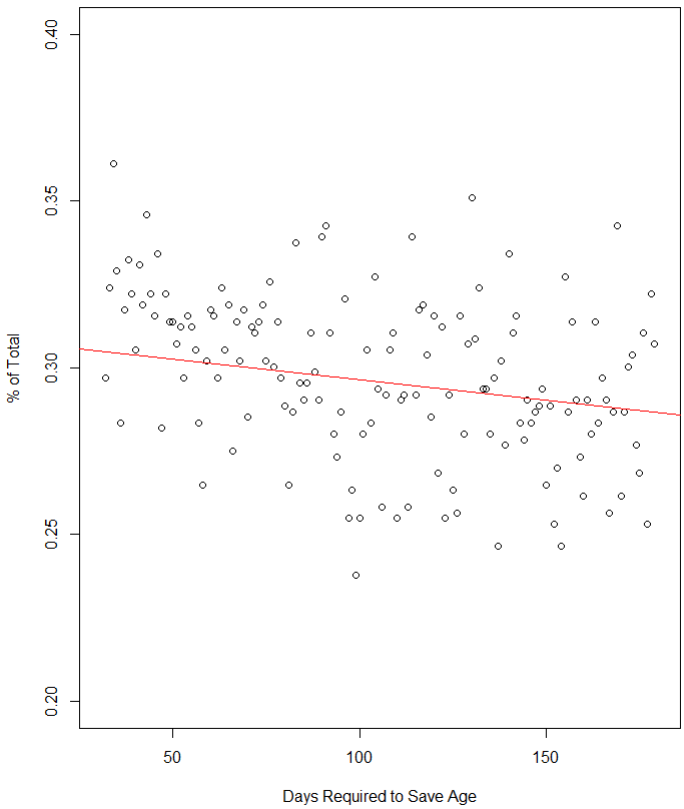
We wish to examine if information about a consumer's lapse risk is contained within consumer's decision to backdate. However, our above model is not able to separately identify an effect from the decision to backdate or an effect from simply having backdated. In other words, we are unable to tell with this initial specification if differential lapse behavior is due to selection based on prior knowledge of lapse risk or if it is based on the differential pricing structure caused by backdating. Since we control for differences in premiums, any non-selection effect must be due to the initial tariff. This tariff is only paid at the start of the policy, thus consumers who act differently based only on having paid the tariff (not selection effects) must be falling prey to a classic sunk cost fallacy.

We are concerned with the selection effect and sunk cost effect separately, rather than the net effect. Thus our identification problem is classic selection-bias which we address through instrumental variables. We instrument for $\mathbf{Backdate_i}$ using the difference between the approval date of the policy and the consumer's birthday ($\mathbf{BackdateOpp_i}$), i.e., how large of an initial tariff the consumer must pay in order to acquire a lower premium. The exclusion restriction is satisfied assuming that consumers do not consider their birthdays when applying for life insurance. Because consumers are more likely to purchase life insurance close to their birthdays, we drop those who apply to purchase life insurance within a month of their birthdays (both before and after). Figure 2 shows the distribution of days required to save age along with a fitted quadratic curve. The slope

is statistically significantly negative, however this effect is exceptionally small: a consumer is .018% more likely to apply 31 days after their birthday vs. 180 days.[22]

Figure 3

Plot of Days Required to Save Age vs. Percent of Total Sample.



Note: This figure shows the distribution of application days relative to birthdays, i.e. how many days of premiums consumers must pay to backdate. Consumers who purchase within 30 days before or 181 days after) of their birthday are not included.

To instrument here we redefine the error term

$$u_i = exp(\alpha v_i + e_i)$$

$$e_i \sim N(0, \sigma^2)$$

$$Backdate_i = f(Z_i) + v_i$$

$$Z_i = (\mathbf{BackdateOpp_i}, IssueAge_i, Male_i, FaceAmt_i, AnnPrem_i,$$

[22]This can be attributed to *Benford's Law* mod *365*.

$$UWClass_i, TermLength_i, IssueYear_i, State_i)$$

Where $f(Z_i)$ is fitted via a Probit specification.

This structure allows us to use a control function approach.[23] Our identified model is

Full Model:

$$\lambda(Lapse_{it}|X) = \lambda_0(t)exp(X_i\beta')u_i \tag{6}$$

$$X_i = (\mathbf{Backdate_i}, IssueAge_i, Male_i, FaceAmt_i, AnnPrem_i, UWClass_i, TermLength_i, IssueYear_i, State_i)$$

$$u_i = exp(\alpha v_i + e_i) \tag{7}$$

$$e_i \sim N(0, \sigma^2) \tag{8}$$

First Stage:

$$\mathbf{Backdate_i} = f(Z_i) + v_i \tag{9}$$

$$Z_i = (\mathbf{BackdateOpp_i}, IssueAge_i, Male_i, FaceAmt_i, AnnPrem_i, UWClass_i, TermLength_i, IssueYear_i, State_i)$$

Second Stage:

$$\lambda(Lapse_{it}|\hat{X}) = \lambda_0(t)exp(\hat{X}_i\beta')e_i \tag{10}$$

$$\hat{X}_i = (\hat{\mathbf{v}}_\mathbf{i}, \mathbf{Backdate_i}, IssueAge_i, Male_i, FaceAmt_i, AnnPrem_i, UWClass_i, TermLength_i, IssueYear_i, State_i).$$

Here the second stage (10) is the same as (5) with the inclusion of the residuals from the first stage (9).[24] Following Basu and Coe (2015) we use Anscombe residuals (Anscombe and Tukey, 1963) to account for the relative rarity of backdating.[25] Equation (10) allows us to separately identify the selection effect (coefficient on $\hat{\mathbf{v}}_\mathbf{i}$) from the sunk cost effect (coefficient on $\mathbf{Backdate_i}$).

[23] Also commonly referred to as two-stage residual inclusion. This method originated with Heckman and Hotz (1989). For an excellent review with a health focus, the reader is referred to Terza et al. (2008).

[24] It is important to note that $-1 < v_i < 1$ and thus is v_i is not normally distributed. However, $E[v_i] = 0$ since the predictions of a Probit specification are unbiased. In Equation (10), v_i is transformed into approximately normal $\hat{\mathbf{v}}_\mathbf{i}$ via an Anscombe transformation. See Basu and Coe (2015) for further discussion on 2SRI with a binary first-stage as well as Anscombe residuals.

[25] Anscombe residuals are a transformation of the standard residuals into an approximately standard normal distribution. Standard errors are bootstrapped.

The interpretation of the coefficient on $\hat{\mathbf{v}}_\mathbf{i}$ is similar in intuition to the coefficient on the inverse Mills ratio in the classic Heckman selection model. That is, the significance of the coefficient shows whether some form of selection is occurring. The interpretation of the coefficient on $\mathbf{Backdate_i}$ as only the effect of paying the initial tariff on future decisions depends vitally on the implicit assumption that premium changes are being effectively controlled for via the additively linear term. There is no reason to believe this is true. To account for arbitrary non-linearity in the control variables while still preserving the ability to instrument we turn to a Lasso technique.

3.4. Lasso

The Lasso (least absolute shrinkage and selection operator) is a model selection technique originally developed by Tibshirani (1996) as an improvement on step-wise regression and adapted to Cox hazard models by Tibshirani et al. (1997). The technique is currently popular in the machine learning literature and was introduced to the econometrics literature by Belloni et al. (2012).[26] The Lasso is in the class of l_1-penalized methods of model selection.[27]

The Lasso is beneficial here in two specific ways. First, the model selection allows us to account for (nearly) arbitrary non-linearity in our control variables via polynomial approximation. Rather than including only linear representations of our control variables, we allow linear, squared, and cubic terms as well as all possible two-variable (inclusive of squared and cubed variables) interaction terms and then allow the Lasso to select the important terms. The second benefit of this method of approximating non-linearity in control variables is the preservation of the linear nature of the treatment variables (here $\mathbf{Backdate_i}$ and the selection effect $\hat{\mathbf{v}}_\mathbf{i}$) allowing for control function instrumentation. Formally, our final model is

Final Model:

$$\lambda(Lapse_{it}|X) = \lambda_0(t)exp(X_i\beta')u_i \tag{11}$$

$$X_i = (\mathbf{Backdate_i}, IssueAge_i, Male_i, FaceAmt_i, AnnPrem_i, UWClass_i, TermLength_i, IssueYear_i, State_i)$$

$$u_i = exp(\alpha v_i + e_i) \tag{12}$$

[26] See also Belloni et al. (2014b), Belloni et al. (2014a), Belloni et al. (2016), and Chernozhukov et al. (2015).

[27] That is, they generally take the form: $\beta_{Lasso} = \underset{\beta}{\text{argmin}} \left\{\sum(y - X\beta)^2\right\}$ subject to $\sum|\beta| \leq l_1$.

$$e_i \sim N(0, \sigma^2) \tag{13}$$

First Stage:

$$\mathbf{Backdate_i} = f(Z_i) + v_i \tag{14}$$

$$Z_i = (\mathbf{BackdateOpp_i}, IssueAge_i, Male_i, FaceAmt_i, AnnPrem_i,$$
$$UWClass_i, TermLength_i, IssueYear_i, State_i)$$

Second Stage:

$$\lambda(Lapse_{it}|\hat{L}) = \lambda_0(t)exp(\hat{L}_i\beta'_{Lasso})e_i$$

$$\text{where } \beta'_{Lasso} = \underset{\beta}{\text{argmax}}\,\{l_{Cox}(\beta L)\} \text{ subject to } |\beta| \leq l_1 \tag{15}$$

$$\hat{L}_i = S(\hat{\mathbf{v}}_\mathbf{i}, \mathbf{Backdate_i}; \hat{\mathbf{X}}_\mathbf{i}, \hat{\mathbf{X}}_\mathbf{i}^\mathbf{2}, \hat{\mathbf{X}}_\mathbf{i}^\mathbf{3}, \hat{\mathbf{X}}_\mathbf{i} : \hat{\mathbf{X}}_\mathbf{i}, \hat{\mathbf{X}}_\mathbf{i} : \hat{\mathbf{X}}_\mathbf{i}^\mathbf{2}, \hat{\mathbf{X}}_\mathbf{i} : \hat{\mathbf{X}}_\mathbf{i}^\mathbf{3})$$

Third Stage:

$$\lambda(Lapse_{it}|\hat{L}^p) = \lambda_0(t)exp(\hat{L}_i^p\beta')e_i \tag{16}$$

$$\hat{L}_i^p = (\hat{\mathbf{v}}_\mathbf{i}, \mathbf{Backdate_i}; \hat{L}_i \text{ such that } \beta_{Lasso,i} \neq 0)..$$

Where the first stage is the same as above.

$$\hat{L}_i = S(\hat{\mathbf{v}}_\mathbf{i}, \mathbf{Backdate_i}; \hat{\mathbf{X}}_\mathbf{i}, \hat{\mathbf{X}}_\mathbf{i}^\mathbf{2}, \hat{\mathbf{X}}_\mathbf{i}^\mathbf{3}, \hat{\mathbf{X}}_\mathbf{i} : \hat{\mathbf{X}}_\mathbf{i}, \hat{\mathbf{X}}_\mathbf{i} : \hat{\mathbf{X}}_\mathbf{i}^\mathbf{2}, \hat{\mathbf{X}}_\mathbf{i} : \hat{\mathbf{X}}_\mathbf{i}^\mathbf{3})$$

is the standardized collection of our two variables of interest, all control variables, squared terms, cubed terms, and two-variable interactions. Equation (15) describes the Lasso procedure where $l_{Cox}(\beta L)$ is the likelihood function for the Cox proportional hazard model.[28] Following Belloni et al. (2016) in Equation (16) we estimate the unpenalized Cox proportional hazard using the two treatment variables and all of the control variables whose coefficients were non-zero in the second stage. The included variables and combinations of variables in $\hat{L}_i^p$ can be interpreted as the optimal polynomial form of the control variables that can be represented in a limited (via the choice of l_1) number of terms.

[28] In Equation (15), the coefficients on $\hat{\mathbf{v}}_\mathbf{i}$ and $\mathbf{Backdate_i}$ are not penalized. l_1 is determined via cross-validation techniques (Goeman, 2010). Additionally, the linear versions of all variables in $\hat{\mathbf{X}}_i$ are included in $\hat{L}_i^P$ regardless of their coefficient in the Lasso stage.

4. Results

Table 2: Cox Proportional Hazard Model

	Dependent variable:
	Hazard(Lapse)
Variable of Interest:	
Backdate	-0.285^{***}
	(0.036)
Issue Age	-0.030^{***}
	(0.001)
Controls:	
Male	0.034
	(0.024)
Face Amount($000s)	-0.001^{***}
	(0.000)
Annual Premium	0.000^{***}
	(0.000)
Length of Term 15	-0.368^{***}
	(0.053)
Length of Term 20	-0.240^{***}
	(0.028)
Length of Term 25	-0.353^{***}
	(0.126)
Length of Term 30	-0.388^{***}
	(0.037)
UW Rank2	0.164^{***}
	(0.056)
UW Rank3	0.509^{***}
	(0.101)
UW Rank4	1.141^{***}
	(0.056)
UW Rank5	0.717^{***}
	(0.049)
State Effects?	Yes
Year Effects?	Yes
Observations	24,661
R^2	0.100
Max. Possible R^2	0.997
Log Likelihood	-70,432.380
Wald Test	$2{,}531.580^{***}$
LR Test	$2{,}604.569^{***}$
Score (Logrank) Test	$2{,}648.117^{***}$
Note:	*p<0.1; **p<0.05; ***p<0.01 This table shows equation (5) fitted via the Cox proportional hazards technique. Std. errors are in parentheses.

The results from our initial specification (Equation (5)) are presented in Table 2. The initial regression results are congruent with our theoretical predictions. Policy owners who choose to backdate their life insurance contracts (effectively paying a two-part tariff) signal their lower likelihood

for lapsing by their willingness to pay for time that already has elapsed that only results in net saving if the policy is held for a relatively long period of time. However, this coefficient does not fully identify the selection effect that we seek. Potentially, consumers who backdate may have no additional knowledge of their propensity for lapsing and are instead exhibiting sunk cost fallacy.

Table 3: First Stage Regression

	Dependent variable:
	Backdate
Instrument:	
Backdate Opp	-0.021^{***}
	(0.000)
Issue Age	0.019^{***}
	(0.001)
Controls:	
Male	0.145^{***}
	(0.023)
Face Amount($000s)	0.001^{***}
	(0.000)
Annual Premium	0.000^{***}
	(0.000)
Length of Term 15	0.227^{***}
	(0.044)
Length of Term 20	0.025
	(0.029)
Length of Term 25	0.231^{***}
	(0.087)
Length of Term 30	0.232^{***}
	(0.037)
UW Rank2	0.039
	(0.037)
UW Rank3	-0.082
	(0.085)
UW Rank4	-0.212^{***}
	(0.052)
UW Rank5	-0.097^{***}
	(0.035)
Constant	-0.516^{***}
	(0.088)
State Effects?	Yes
Year Effects?	Yes

Note: *p<0.1; **p<0.05; ***p<0.01
This table shows fitted values for the first stage in our control function procedure (equation (9)). Std. errors are in parentheses.

To identify these separate effects, we exploit inherent randomness in the time of year, relative to the consumer's birthday, that the consumer applies for life insurance. This allows us to perform a pseudo-random experiment exploiting variation in the initial tariff consumers have to pay while

holding constant the reduction in future premiums. The results from the first stage in our control function approach (Equation (9)) are presented in Table 3. Our instrument is strong and loads in the predicted manner – people who have to pay a higher initial tariff, *ceteris paribus*, are less likely to choose to backdate.

Table 4: Two-Stage Residual Inclusion Estimation

	Dependent variable:	
	Hazard(Lapse)	
	Initial	Instrumented
	(1)	(2)
Variables of Interest:		
Backdate	-0.285^{***}	-0.120^{*}
	(0.036)	(0.067)
Stage 1 Residuals		-0.113^{***}
		(0.038)
Controls:		
Issue Age	-0.030^{***}	-0.031^{***}
	(0.001)	(0.001)
Male	0.034	0.028
	(0.024)	(0.024)
Face Amount($000s)	-0.001^{***}	-0.001^{***}
	(0.000)	(0.000)
Annual Premium	0.000^{***}	0.000^{***}
	(0.000)	(0.000)
Length of Term 15	-0.368^{***}	-0.376^{***}
	(0.053)	(0.054)
Length of Term 20	-0.240^{***}	-0.241^{***}
	(0.028)	(0.029)
Length of Term 25	-0.353^{***}	-0.357^{***}
	(0.126)	(0.122)
Length of Term 30	-0.388^{***}	-0.394^{***}
	(0.037)	(0.038)
UW Rank2	0.164^{***}	0.163^{***}
	(0.056)	(0.055)
UW Rank3	0.509^{***}	0.516^{***}
	(0.101)	(0.106)
UW Rank4	1.141^{***}	1.149^{***}
	(0.056)	(0.059)
UW Rank5	0.717^{***}	0.721^{***}
	(0.049)	(0.050)
State Effects?	Yes	Yes
Year Effects?	Yes	Yes
Observations	24,661	24,661
R^2	0.100	0.101
Max. Possible R^2	0.997	0.997
Log Likelihood	-70,432.380	-70,427.620
Wald Test	$2{,}531.580^{***}$	$2{,}545.240^{***}$
LR Test	$2{,}604.569^{***}$	$2{,}614.074^{***}$
Score (Logrank) Test	$2{,}648.117^{***}$	$2{,}661.391^{***}$

Note: *p<0.1; **p<0.05; ***p<0.01
This table presents again for comparison our initial model from Table 2 in the first column and shows fitted values for the second stage in our control function approach in the second column. Bootstrapped std. errors are in parentheses.

We then re-estimate the hazard specification from Table 2 (Equation (5)), this time including the transformed residuals from the first stage. The results from our instrumented model (Equation (10)) are presented in the second column of Table 4. Our original estimated effect was indeed a combination of both selection and sunk cost fallacy. The selection effect (-0.062) is statistically significant, and it also appears that a significant portion (-0.162) of the reduction in lapse likelihood is due to sunk cost fallacy.

The third column of Table 5 shows the results of the Lasso procedure. In addition to the 12 independent variables shown in the table, the model includes state effects, year effects, and 135 (out of over 800 potential) other forms of the control variables (either squared, cubed, or interaction) that the Lasso procedure selected as important.[29] The inclusion of these variables does not change the significance of our main results, though the point estimates are slightly diminished.[30] The selection effect does not have an interpretation beyond directional comparative statics, however the sunk cost effect does. Exponentiating the coefficient on $\mathbf{Backdate_i}$ results in a marginal relative hazard rate of 89.2%. That is, the sunk cost effect of backdating reduces the per-period hazard rate of lapsing by 10.8%.

[29]The extra control variables are suppressed in the table for space; a full table reporting the coefficients on all Lasso control variables can be found on the author's website.

[30]Interestingly, the significance of the coefficients for the linear terms of $Male$ and $FaceAmount$ goes away. This is due to the inclusion, via the Lasso, of many significant interaction terms containing those variables.

Table 5: Lasso Estimation

	Dependent variable:		
	Hazard(Lapse)		
	Initial	Instrumented	Lasso Controlled
	(1)	(2)	(3)
Variables of Interest:			
Backdate	-0.285^{***}	-0.120^{*}	-0.114^{*}
	(0.036)	(0.067)	(0.062)
Stage 1 Residuals		-0.113^{***}	-0.067^{*}
		(0.038)	(0.037)
Controls:			
Issue Age	-0.030^{***}	-0.031^{***}	-0.031^{***}
	(0.001)	(0.001)	(0.004)
Male	0.034	0.028	-0.170
	(0.024)	(0.024)	(0.111)
Face Amount($000s)	-0.001^{***}	-0.001^{***}	0.001
	(0.000)	(0.000)	(0.001)
Annual Premium	0.000^{***}	0.000^{***}	-0.002
	(0.000)	(0.000)	(0.004)
Length of Term 15	-0.368^{***}	-0.376^{***}	-0.218
	(0.053)	(0.054)	(0.248)
Length of Term 20	-0.240^{***}	-0.241^{***}	-0.561^{***}
	(0.028)	(0.029)	(0.123)
Length of Term 25	-0.353^{***}	-0.357^{***}	-0.050
	(0.126)	(0.122)	(0.308)
Length of Term 30	-0.388^{***}	-0.394^{***}	-0.284^{**}
	(0.037)	(0.038)	(0.129)
UW Rank2	0.164^{***}	0.163^{***}	0.284^{*}
	(0.056)	(0.055)	(0.171)
UW Rank3	0.509^{***}	0.516^{***}	0.794^{**}
	(0.101)	(0.106)	(0.361)
UW Rank4	1.141^{***}	1.149^{***}	0.885^{**}
	(0.056)	(0.059)	(0.348)
UW Rank5	0.717^{***}	0.721^{***}	0.471^{*}
	(0.049)	(0.050)	(0.282)
State Effects?	Yes	Yes	Yes
Year Effects?	Yes	Yes	Yes
Lasso Controls? (Number)	No	No	No
Observations	24,661	24,661	24,661
R^2	0.100	0.101	0.126
Max. Possible R^2	0.997	0.997	0.997
Log Likelihood	-70,432.380	-70,427.620	-70,075.100
Wald Test	$2{,}531.580^{***}$	$2{,}545.240^{***}$	$2{,}408.640^{***}$
LR Test	$2{,}604.569^{***}$	$2{,}614.074^{***}$	$3{,}319.120^{***}$
Score (Logrank) Test	$2{,}648.117^{***}$	$2{,}661.391^{***}$	$3{,}594.168^{***}$

Note: *p<0.1; **p<0.05; ***p<0.01
This table presents again for comparison both our initial model from Table 2 in the first column and the fitted coefficients for the second stage in our control function approach in the second column. The third column represents the results of the Lasso procedure. A full representation and list of the Lasso coefficients can be found on the author's website. Bootstrapped std. errors are in parentheses.

5. Conclusions

Our results indicate that asymmetric information about lapse risk can be reduced through a firm offering a menu of two-part tariff contracts and allowing consumers to self-select into the contract designed for their lapse type. These optional two-part tariffs serve as a screening device for insurers on consumers' likelihood of lapsing. Consumers who choose to pay the two-part tariff (e.g., backdating their life insurance contracts) signal their lower likelihood for lapsing by their willingness to pay for time that already has elapsed to have lower premiums that only results in net savings if the policy is held for a long time. If consumers terminate the policy early, they do not reap the benefit of the lower premium level. Such a willingness translates into a significantly lower hazard of lapsing, thus aligning the interests of the consumer with the insurer. Our research provides key insight into why insurers do not use continuous (with regards to age) pricing despite the computational ease of doing so – the value of the screening provided by offering the optional two-part tariff outweighs any actuarial downside caused by discreteness in years.

We additionally find strong evidence that life insurance consumers exhibit behavior consistent with sunk cost fallacy in their lapsing behavior, even when controlling for arbitrary non-linearity in premium effects on lapse proclivity. This interesting finding implies a larger degree of reverse causality with lapse rates and premium structures (including initial tariffs) than is currently being discussed in the literature. Development of a simultaneous model of lapse rates and premium structures would be an excellent direction for further research. While backdating is an example of an optional two-part tariff, it is not the optimal one. The value of the screening mechanism varies randomly from consumer to consumer with the randomness of the distance to their birthday – someone signing up the day after their birthday is almost sure to backdate and thus there is very little selection effect to signal. Additionally, by definition, backdating is a binary choice. An optimal menu of contracts would likely consist of the consumer choosing what portion of their underwriting cost to pre-pay and having a subsequent reduction in premiums reflective of their choice. Further examination of this optimal structure is worthy of additional study.

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