The Effect of Life Insurance Policy Provisions on Suicide Rates

Samuel Hsin-yu Tseng*

University of Chicago

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Abstract

The economics of asymmetric information has been well developed in theoretical terms, yet little empirical work has been done to apply it to life insurance markets. To address this issue, I investigate the incentive effects of a two-year suicide exclusion in individual life insurance policies. The findings suggest that the suicide rate quadruples after the exclusion period and that suicides during the exclusion period are often disguised as accidental deaths. The results match the implications of the theory and suggest that economic considerations affect the timing and method of committing suicide.

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1. Introduction

Since the seminal contributions of Akerlof (1970), Spence and Zeckhauser (1971), and Rothschild and Stiglitz (1976), insurance markets have been employed as a classic example of asymmetric information. On the one hand, numerous theoretical models have been constructed to explore the implications of asymmetric information[1]. On the other hand, a variety of empirical studies have tested these models[2]. As a survey paper**,** Chiappori and Salonie (2003) find mixed evidence of asymmetric information in insurance markets. Specifically, many authors have found evidence of adverse selection in annuity markets[3]; however, most studies of life insurance markets have found little or no evidence of asymmetric information[4].

To address this issue, I investigate the incentive effects of a two-year suicide exclusion. In the United States, all individual life insurance policies include a suicide exclusion, which states that no death benefits will be paid if the insured dies by suicide within an exclusion period, usually the first two policy years. This exclusion indicates that suicide becomes a covered risk after this period. Since this exclusion creates a credible exogenous change in payable death benefits, it constitutes a natural experiment on assessing the incentive effects that it causes and consequently allows for testing the asymmetric information hypothesis in life insurance markets.

The presence of asymmetric information suggests that this suicide exclusion should affect the insureds' incentive to commit suicide and therefore lead to two testable hypotheses. The first hypothesis suggests an abnormal increase in suicide rates directly after the exclusion period. This abnormal increase is caused by two types of moral hazard. The first type concerns pure moral hazard, which assumes that ex ante none of the insureds have suicide intentions, yet ex post some of the insureds suffer traumatic events that lead them to consider committing suicide. Since insured persons have control over the timing of their suicides, they may defer committing suicide

[1] See Dionne (2000) for a comprehensive review.
[2] For recent empirical studies on automobile insurance, see Chiappori and Salonie (2000), Cohen (2002), and Cohen and Dehejia (2003); for studies on health insurance, see Culter (2002) for a review and Cardon and Hendel (2001).
[3] These studies include Carlson and Lord (1986), Friedman and Warshawsky (1990), Bruggiavini (1993), and Finkelstein and Poterba (2004).
[4] Prior empirical studies on life insurance include Beliveau (1984) and Cawley and Philipson (1999), yet neither of them finds evidence of asymmetric information. Beliveau (1984) examines the relationship between premium rates and amounts of coverage, and her findings match the prediction of Rothschild and Stiglitz (1976). However, her analysis lacks crucial variables, such as smoker status, observed by the insurer. Consequently, her findings provide insufficient evidence of asymmetric information (Villeneuve, 2000). Cawley and Philipson (1999) examine the relationship between price and quantity and the covariance between contract size and risk in the life insurance market. Their findings suggest bulk discounts, contrary to convex pricing, and suggest that low-risk individuals carry more insurance coverage. Therefore, their study finds no evidence of asymmetric information.

until it becomes a covered risk. The second type combines adverse selection and moral hazard, which assumes that ex ante some of the insureds have sustainable suicide intentions, and they choose to commit suicide after it becomes a covered risk[5].

The second testable hypothesis concerns "disguised suicides" and predicts a decline in rates of accidental deaths after the exclusion period[6]. During the exclusion period, life insurance policies pay death benefits to all causes of death except suicide. Therefore, to secure death benefits during this period, the insureds who intend to commit suicide have the incentive to cleverly disguise their suicidal behaviors so that their deaths could be relabeled as other causes of death, such as accidental deaths[7].

To test the aforementioned predictions, I acquired a mortality data set from the Society of Actuaries (SOA). This data set includes mortality records of a sample of people who owned individual life insurance policies classified as preferred or standard risks and who died in the period of 1990-1995[8]. Analysis of these data confirms both of the aforementioned hypotheses and suggests that the suicide exclusion has significant incentive effects on the timing and method of committing suicide. Figure 1 shows that the suicide rate substantially increases after the second policy year; this increase does not simply reflect increases in general mortality because Figure 2 shows a small rise in natural death rates after the second policy year. The regression analyses suggest that controlling for appropriate factors, the suicide rate quadruples after the exclusion period. Moreover, by using a "difference in difference" approach, I find that disguised suicides account for 34% of the overall suicides resulting from the implementation of the suicide

[5] I attempted to distinguish these two types of moral hazard. However, the limited data yielded insignificant results. In theory, this distinction can be made by comparing the ratio between the suicide rate during the third policy year and the suicide rate during the fourth policy year. When a two-year suicide exclusion is included, and traumatic events are uniformly distributed across years, the pure moral hazard suggests that the suicide rate should be the highest during the third policy year, decline by two-thirds during the fourth year, and remain the same from the fourth year on. On the contrary, the second type, which combines adverse selection and moral hazard, suggests that the suicide rate should resemble a spike with a positive rate during the third policy year, but zero in the other years. See Section 6 for further discussion.

[6] This type of moral hazard is facilitated by the difficulty in verifying suicide and may be facilitated by the preference of coroners or medical examiners. When death claims due to cleverly disguised suicides are filed, insurance companies are unable to deny such claims unless they can verify that these deaths are caused by suicide. In practice, it is very difficult to verify suicide because this requires a suicide note to be discovered or that the death cannot be reasonably explained in any other way. Additionally, when causes of death are ambiguous, coroners or medical examiners acquainted with the beneficiaries may prefer to classify the causes as accidental deaths rather than suicides.

[7] According to Phillips and Ruth (1993), suicides are possibly relabeled as three categories of accidental deaths: automobile accidents, pedestrian deaths, and accidental barbiturate poisoning.

[8] Since the SOA data set only includes information of the insured, it is a truncated sample. This suggests that my analysis applies to the pool of the insured, but not to all members of the population.

exclusion.

While researchers have been trying to understand suicide from its clinical, neurobiological, legal, and psychosocial aspects, this paper attempts to understand suicide from the economic perspective by investigating the incentive effect of the suicide exclusion. This exclusion is relevant because the death benefit payment, whose amount can reach $1 million, is not a small amount of money. Studies suggest that economic factors affect suicide rates; the findings of Hamermesh and Soss (1974) and Culter, Glaeser, and Norberg (2001) suggest that income level is negatively correlated with suicide rates[9]. Moreover, this paper matches the studies on suicidology. Durkheim (1951) uses altruism to explain suicide rates; this altruistic suicide is confirmed by a quadrupled suicide rate directly after the exclusion period.

As is well known, one of the major difficulties in exploring evidence of asymmetric information is to control selection problems. This difficulty is avoided by two features of this paper even though this paper does involve a greater level of aggregation than previous studies. First, the suicide exclusion creates an exogenous change in death benefit payments; this makes it easier to overcome selection problems because no other factors have drastically changed at the time when the suicide exclusion expires. Thus the resulting change in suicide behavior can safely be attributed to the incentive effect caused by the suicide exclusion. Secondly, the investigation of this incentive effect requires a sample of the insured only; this mitigates selection problems because the screening capability of the underwriting process suggests that the people having a history of mental or physical health impairments are excluded from the sample.

The rest of the paper is structured as follows. Section 2 overviews the suicide exclusion. Section 3 describes the data and compares suicide rates between the general population and the insured. Section 4 estimates the incentive effect of the suicide exclusion. Section 5 discusses estimation bias and research extensions. Section 6 constructs theoretical models to illustrate the two types of moral hazard: (1) pure moral hazard; (2) moral hazard with adverse selection. Section 7 concludes.

[9] Hamermesh and Soss (1974) analyze the relationship between permanent income and suicide by employing time-series and cross-sectional data on male suicide rates. Their findings suggest that for male adults aged 25 and above, permanent income level is negatively correlated with suicide rates. Culter, Glaeser, and Norberg (2001) analyze the rising suicide rate among youths in 1950-1990. Their findings suggest that female divorce rate contributes most to the rising suicide rate and that counties with a higher income level have a lower youth suicide rate.

2. Overview of the Suicide Exclusion

A suicide exclusion is included in individual life insurance policies so that two competing interests can be balanced. On the one hand, this exclusion mitigates the problem of adverse selection because it deters applicants who intend to secure benefits for their beneficiaries by committing suicide shortly after their policies are issued. On the other hand, by covering suicide after the exclusion period, this exclusion fulfills the primary function of life insurance that financially protects all of the beneficiaries. When the insureds die by suicide, the resulting economic loss to their beneficiaries is the same as the loss if they had died by other causes. This exclusion has been a standard provision in individual life insurance policies at least since 1950 (Anderson, 1951).

The length of the exclusion is affected by state regulations and company practices. Thirty-three states[10] allow for a two-year suicide exclusion, Colorado and North Dakota allow for a one-year exclusion, but Missouri imposes a more stringent regulation, requiring that the insurer cannot use suicide to deny death benefit payments unless it can be shown that the insureds intend to commit suicide when they applied for the policies[11]. Moreover, although fourteen states[12] and the District of Columbia do not specifically address the issue of suicide exclusion in their statutes, some of these states, such as Oregon, permit a two-year suicide exclusion in their administrative rules.

Additionally, I surveyed the twenty insurance companies contributing to the SOA mortality data set and found that the standard practice is to include a two-year suicide exclusion on policies sold in all states except Colorado and North Dakota. In these two states, a one-year exclusion is included. This standard practice is adopted by all companies except one, which includes a one-year suicide exclusion for all states. Finally, among all twenty companies, the length of suicide exclusion does not vary by product types.

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[10] These thirty-three states include Alabama, Alaska, Arizona, Arkansas, California, Delaware, Georgia, Hawaii, Idaho, Illinois, Kentucky, Louisiana, Maine, Maryland, Minnesota, Montana, Nebraska, Nevada, New Hampshire, New Mexico, New York, North Carolina, Oklahoma, South Carolina, South Dakota, Tennessee, Texas, Utah, Virginia, Washington, West Virginia, Wisconsin, and Wyoming.

[11] It is practically impossible for the insurer to verify such an intention, so the Missouri statute has been interpreted as invalidating suicide exclusion.

[12] These fourteen states include Connecticut, Florida, Indiana, Iowa, Kansas, Massachusetts, Michigan, Mississippi, New Jersey, Ohio, Oregon, Pennsylvania, Rhode Island, and Vermont.

3. Data Description and Suicide Rate Comparison between the General Population and the Insured

To help companies conduct mortality comparisons, the Society of Actuaries (SOA) undertakes continuous studies on mortality experience under standard ordinary life insurance. These studies include policies fulfilling five requirements: (1) issued in the United States; (2) marketed by regular channels (e.g., agents, but not mail order); (3) not held jointly nor having family policies or riders; (4) individually underwritten by regular methods (i.e., non-medical, paramedical, or medical examinations); (5) classified as preferred or standard risks. These studies provide ideal data to investigate the incentive effect of the suicide exclusion because they include all of the variables used for pricing mortality risks[13] and the information on death claims[14]. However, due to confidentiality agreements between the SOA and contributing companies, only the aggregate data at the national level have been acquired. These aggregate data come from the studies examining mortality experience in 1990-1995 and include information on issue age, gender, policy duration, cause of death, and death rate. These are the data used in my investigation.

This data source has been rarely used in economic studies. Therefore, it is important to understand the degree to which the data represent the mortality experience of all companies in the United States. The SOA is the only organization that has the authority to collect mortality data from multiple companies, so its data are more representative than those from other sources; however, the SOA mortality data tend to represent larger companies and companies with lower mortality rates. These two aspects of the data and the issue of data accuracies are discussed in the data appendix.

Table 1 summarizes the cause-of-death data in the 1990-1995 SOA mortality studies. The data include 7,130 policies terminated by suicide, which accounts for 1.87% of the policies terminated by deaths. In Table 2, I compare the suicide rate of the general population with that of the insureds within the first ten policy years. This duration is chosen so that the comparison can be conditional on the same age of death. For the general population, the suicide rate is measured by the number of people who die by suicide per 100,000 people. For the insured, however, because each age group’s by-count suicide rate (the number of policies terminated by suicide per

[13] These variables include issue age, gender, smoker status, type of underwriting requirement, risk classification, type of policy, and amount of coverage.
[14] The information includes cause of death, duration, and claim amount.

100,000 policies in force) cannot be acquired, I employ the by-amount suicide rate [the amount of death benefits paid to suicide[15] per $100,000 insurance coverage exposed to loss (exposure)[16]]. Consequently, the suicide rate of the insured can be viewed as the weighted suicide rate of the general population, with the weight being the amount of benefits.

Column (1) of Table 2 shows that for the general population, the male suicide rate remains flat, approximately 24 suicides per 100,000 people, between the age of 15 and 64. The suicide rate of the elderly male aged 65 and above is 1.6 times higher than that of the male aged 15-64. Moreover, the male suicide rate averages 4.3 times higher than that of the female. Quinnett (1997) suggests that this may result from women's willingness to seek professional help for treatment of depression.

Column (2) of Table 2 shows that this gender difference in suicide rates also exists among the insured. However, the suicide rate of the insured substantially differs from that of the general population in two aspects. First, the suicide rates of the insured are approximately one-third of those of the general population. Although this substantial gap may result from the difference in units of suicide rates, this difference alone cannot fully explain the gap because it is likely to be caused by other factors, such as differences in mental health condition, income, and insurance status. The screening capability of the underwriting process suggests that insured adults are mentally healthier than the general population. Insured adults also tend to earn more income than the general population[17]. These two factors suggest that insured adults have a lower suicide rate. However, death benefit payments may lead insured adults to have a higher suicide rate.

Secondly, in the general population, the suicide rate of male youths (aged 15-24) amounts to 89% of the suicide rate of male young adults (aged 25-34). For the insured, however, the suicide rate of male youths is 2.3 times higher than that of male young adults[18]. This difference

[15] Since suicide is an excluded risk during the first two policy years, the amount of coverage owned by the insured committing suicide is employed to calculate suicide rates during this period. This adjustment also applies to other death claims in which no death benefits are paid out.
[16] The term of exposure takes into account the length of time when a policy is in force during a study year. For example, when mortality experience of the calendar year 2003 is studied, the amount of exposure is $100,000 for a policy with coverage $100,000 in force during the whole calendar year. However, the amount of exposure is $50,000 only for a policy with coverage $100,000 in force during July 1, 2003 and December 31, 2003.
[17] By reporting the profile of households owning life insurance in 1992, the American Council of Life Insurance (1994) suggests that income level is positively correlated with the amount of life insurance coverage. For insured households with an annual income of $35,000-$49,999, their amount of coverage averaged $115,600. This amount was $35,600 for insured households with an annual income less than $25,000.
[18] A similar difference also exists in accidental deaths because insured male young adults have a substantially lower

suggests that people who purchase life insurance for their own lives are more cautious about the consequences of their deaths and therefore are less likely to commit suicide. Most young adults purchase life insurance for their own lives, while policies written on the life of youths are usually purchased by their parents or grandparents.

4. Investigations of the Incentive Effect of the Suicide Exclusion

The incentive effect of the two-year suicide exclusion is investigated in two steps. First, I examine the relationship between the suicide rate and policy duration for insured adults whose issue ages are between 20 and 59 and whose policies are within the first eight years[19]. Secondly, to seek evidence of disguised suicides, I examine the relationship between the accidental death rate and policy duration for the same group of insured adults.

Figure 1 shows that the male suicide rate amounts to 0.002% of the exposure during the first two policy years; however, this rate quadruples during the 3rd-4th policy years and stays almost constant after that. Table 3 shows that this pattern holds for every age and gender classification. This suggests that the exclusion affects the timing of suicide among the insured.

However, this effect may be spurious because of the selection effects caused by the underwriting process[20] and by voluntary policy termination. Suicide rates are expected to be higher in later policy years because suicide risk is increased by poor health, and health conditions of the insureds deteriorate as policy duration increases. This deterioration is caused by aging and compounded by the underwriting process and by voluntary policy termination. Table 2 shows that in the general population, suicide rates remain flat for adults dying at the age of 25-64; this suggests that the spurious effect caused by aging is peripheral.

To further investigate the aforementioned selection effects, I look into natural death rates.

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accidental death rate than insured male youths. For the insured, the accidental death rate of male youths amounts to $38.2 death benefits paid to accidental deaths per $100,000 exposure; this rate is $17 for male young adults. In the general population, however, the accidental death rate of male youths amounts to 59.6 accidental deaths per 100,000 people; this rate is 52.5 for male young adults.

[19] I chose this age group for two reasons. First, it excludes juvenile policies issued at the age of 0-19 because the purchase decisions of these policies are not made by the insureds themselves. Secondly, it excludes policies issued for the elderly aged 60 and above because the volume of these policies is small, and the elderly has a different pattern of suicidal behavior. Additionally, this analysis includes policies within only the first eight years in order to exclude the adverse selection effect caused by “select and ultimate term” policies. For overviews of these policies, see Dukes and MacDonald (1980) and SOA (1983; 1984).

[20] Being rated standard or higher in the underwriting process is required to purchase new policies; this suggests that unhealthy insureds are more likely to retain their policies.

Figure 2 shows that the natural death rate slightly increases after the second policy year; the male natural death rate during the 3rd-4th policy years is only 1.4 times higher than that during the first two policy years. In the data appendix, I calculate the upper bound value of the selection effects[21]; this calculation suggests that the male natural death rate during the 3rd-4th policy years is at most 1.7 times higher than that during the first two policy years**.** Thus both the figure and the calculation suggest relatively small selection effects.

To estimate the incentive effect of the suicide exclusion, I regress suicide rates on indicator variables of policy duration and relevant covariates. In the estimation, heteroscedasticity may arise and lead to inefficient estimators of ordinary least squares. To correct for heteroscedasticity, I employ the method of weighted least squares (WLS) in which the weight is the distribution of exposure. The observations come from the grouped data in Table 3, so the unit of observation is a combination of issue age and policy duration. Table 1 indicates that these groups have unequal amounts of exposure.

In summary, I assume that the suicide rate depends on four factors: gender, suicide exclusion, issue age, and selection effects caused by the underwriting process and by voluntary policy termination. The effects of the latter two factors are proxied by natural death rates, so the estimation is performed as follows:

(1) $\ln (S)_{it} = \alpha + \beta (DA)_{it} + \gamma (FD)_{it} + \lambda \ln (RND)_{it} + \varepsilon_{it}$

where S_{it} represents the suicide rate of the group with issue age *i* and in policy year *t*; DA represents a duration dummy, which equals 1 for duration after the second policy year; FD represents a female dummy; RND represents the natural death rate. The coefficient of DA (β) captures the incentive effect of the two-year suicide exclusion; if the estimate of β is positive and significant, this suggests that the suicide rate increases after the second policy year and therefore suggests evidence of asymmetric information.

The results are reported in Table 4. In Column (1), the method of ordinary least squares generates a statistically significant positive estimate of the duration dummy, but the estimate of the natural death rate is insignificant. This insignificance may result from the presence of heteroscedasticity because the estimated standard errors are reduced in Column (2) employing

[21] This upper bound value assumes that both the healthy and unhealthy insureds exist during the 1st-2nd policy years, while only the unhealthy insureds exist during the 7th-8th policy years. The unhealthy insureds retain their policies, while the healthy insureds lapse their policies over time. The rate of lapsation is assumed to be constant.

the method of weighted least squares. Specifically, the estimate of the duration dummy is significant at the 1 % level with a magnitude of 1.45. This suggests that the suicide rate averages 4 (i.e., $e^{1.45}$) times higher after the second policy year. Column (3)-(6), which include each gender separately, also suggest that the suicide rate significantly increases after the exclusion period.

Table 5, which reports accidental death rates across policy years, suggests that the suicide exclusion causes the insureds to disguise their suicides as accidental deaths during the 1^{st}-2^{nd} policy years. Compared with the first two policy years, accidental death rates decline during the 3^{rd}-4^{th} policy years, but do not decline further during the 5^{th}-6^{th} policy years among male adults with issue ages of 30-59 and among female adults with issue ages of 20-59.

However, this decline in accidental death rates may be spurious due to two competing effects: age effects and selection effects caused by the underwriting process and by voluntary policy termination. As policy duration increases, age effects suggest a decline in accidental death rates[22], while the selection effects suggest an increase. When these two kinds of effects are assumed to cancel each other out, the decline in accidental death rates between the 1^{st}-2^{nd} and 3^{rd}-4^{th} policy years can be attributed to disguised suicides only. This assumption allows me to estimate the magnitude of disguised suicides. Furthermore, by employing a "difference in difference" approach, I compute the magnitude of the overall suicides resulting from the implementation of the exclusion. That is, I compute the differences in death rates between the 1^{st}-2^{nd} and 3^{rd}-4^{th} policy years for suicide and for accidental deaths. Then I take the difference between the difference in suicide rates and the difference in accidental death rates. The results summarized in Table 6 suggest that the male death rate due to disguised suicides amounts to 0.00327% of the exposure. This implies that disguised suicides account for 34% [i.e., $\frac{3.27}{(6.44+3.27)}$] of the overall suicides resulting from the implementation of the suicide exclusion.

Actually, the true percentage of disguised suicides may be more than 34% due to a rising trend in accidental death rates; Table 5 shows that male accidental death rates during the 3^{rd}-4^{th}

[22] Mortality statistics of the general population suggest that accidental death rates decline steadily from the age group of 15-24 to the age group of 45-54. In 1990-1995, the accidental death rate of male youths aged 15-24 amounts to 59.6 accidental deaths per 100,000 people. This rate is 52.5 for male young adults aged 25-34, 47.9 for male adults aged 35-44, and 41.8 for male adults aged 45-54.

policy years are lower than those during the 5th-6th policy years and those during the 7th-8th policy years[23]. This trend implies that the selection effects predominate over age effects.

Finally, Table 6 suggests that disguised suicides are more prevalent among male adults with issue ages of 40-59 compared with young male adults with issue ages of 20-39. This table also suggests that the female has a higher percentage of disguised suicides.

5. Estimation Bias and Extensions

The findings suggest a quadrupling in the suicide rate after the exclusion period, but this magnitude may be affected by estimation bias. I discuss this issue and outline possible research extensions below.

Estimation bias may result from omitted variables correlated with policy duration and from variations in the lengths of the suicide exclusion. Due to the unavailability of data, this paper omits other socioeconomic variables that may be correlated with policy duration[24]. Specifically, it omits the effect of divorce. This omission may lead to an exaggeration of the incentive effects of the suicide exclusion because the suicide risk of divorced men is twice as high as that of married men (Kposowa, 2000), and male insureds may become divorced as policy duration increases.

On the other hand, the variations in the lengths of suicide exclusion lead the findings to underestimate the incentive effect of the two-year suicide exclusion because a one-year exclusion is included on one company's policies and on policies sold in Colorado and North Dakota. Furthermore, the suicide exclusion is practically invalidated in Missouri. This company submitted cause-of-death information, and its policy records account for 31% of the data used in my investigation. Between 1992 and 1995, the number of individual life insurance policies in force in Colorado and North Dakota accounted for 1.5% of the policies in the United States. This percentage was 2.4% for Missouri.

Moreover, the findings are derived from aggregate data at the national level, but studies

[23] I use accidental death rates during the 3rd-4th policy years as the base because disguised suicides may affect the rates during the first two policy years.

[24] Racial groups and geographic locations are two of the omitted variables that affect suicide rates, but seem uncorrelated with policy duration. Whites and Native Americans have higher suicide rates than other racial groups, and suicide rates are higher in the western states. Additionally, the practice of determining data values suggests that there is likely no correlation between policy duration and variables used for pricing mortality risks. The values of all of the variables used for pricing mortality risks are determined at the time when policies are issued and do not change as policy duration increases; the two-year suicide exclusion applies to all types of policies.

suggest that the use of aggregate data may result in estimation bias (e.g., Cutchin and Churchill, 1999; Kunce and Anderson, 2002; McGuckin and Stiroh, 2002). These aforementioned biases can be remedied if the individual policy records mentioned in Section 3 are secured.

Additionally, these individual policy records will allow me to extend this paper by distinguishing several types of asymmetric information. First, I can test for the type of adverse selection that suggests the insureds who have suicide intentions ex ante choose term policies. The summary statistics of the SOA studies show that term policies have a higher suicide rate, 8% of death benefit payments, versus 3% in permanent policies. Since the premium rate of term policies is much lower, about one-fifth of that of permanent policies, this suggests that the higher suicide rate may be caused by the difference in income or by adverse selection. Secondly, I can test for the type of asymmetric information that suggests the insured committing suicide purchases a larger amount of coverage than the average insured whose death is due to other causes. The summary statistics of the SOA studies show that the by-count suicide rate averages 1.87% of the policies terminated by deaths, while the by-amount suicide rate reaches 4.28% of the death benefit payments.

Finally, this paper can be extended to investigate whether death benefit payments lead the insured to have a higher suicide rate than the uninsured. This research can be carried out by using the data of the Health and Retirement Study (HRS), the mortality data provided by the National Center for Health Statistics (NCHS), and the Census data[25]. Since life insurance policies pay out death benefits, the presence of asymmetric information predicts an incentive effect that leads the insured to have a higher suicide rate than the uninsured. However, to explore this incentive effect, selection problems have to be carefully controlled because Table 2 shows that the observed suicide behavior contradicts the aforementioned prediction: the insured actually has a lower suicide rate than the general population.

6. Theoretical Models Characterizing Incentive Effects of the Suicide Exclusion

To characterize the incentive effect of the suicide exclusion, I propose two theoretical models: a pure moral hazard model and a model that combines adverse selection and moral

[25] The HRS data includes variables of life insurance coverage but does not have suicide information. The NCHS mortality data, on the other hand, has suicide information but does not have information on insurance status.

hazard. The pure moral hazard model assumes that ex ante none of the insured has suicide intention. However, ex post some of the insured suffer traumatic events that lead them to consider committing suicide. On the other hand, the model that combines adverse selection and moral hazard assumes that ex ante some of the insured have suicide intention. This section is structured as follows. First, I illustrate the setup of households and insurance markets that applies to both models. Second, I analyze each model separately. Finally, I highlight the testable implications of both models.

6.1 Households and Insurance Markets

Consider an economy inhabited by many households whose number is normalized to 1. Members of the households live for T periods. They will not die prematurely unless they commit suicide. In addition, the head of each household is endowed with one unit of labor used to produce output. Household heads care about their own consumption and the bequest left to their beneficiaries, who are the rest of the members in the household. Thus, household heads use their output to purchase two types of goods: consumption and life insurance. Furthermore, household heads can leave bequests to their beneficiaries only through life insurance because other channels do not exist. For the sake of tractability, I assume that in each household, only the head is eligible to purchase life insurance. In addition, household heads are risk averse and have concave utility functions over consumption and death benefits received by their beneficiaries. Formally, for insured household heads who die in period T, their utility function is

$$U = \sum_{t=1}^{t=T} \beta^{t-1} u(c_t) + \beta^T u(Benefits).$$

Regarding insurance markets, I assume that insurance companies face Bertrand competition. That is, in period 0, companies simultaneously offer identical life insurance policies, set their premiums, and let the market determine the amount of insurance coverage sold. Thus, in equilibrium, premiums will equal marginal cost, and each company will earn zero profits[26]. These identical insurance policies have the following four features. First, these policies pay death benefits, worth B amount of output, to the insured's beneficiary in the period right after the insured's death. Secondly, these policies include a suicide exclusion, which states that if the insured dies by suicide in period 1, no death benefits will be paid. Thirdly, exclusivity applies:

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[26] In the presence of asymmetric information, insurance companies in competitive markets may earn positive profits if any of the following three conditions is fulfilled: (1) non-exclusivity; (2) the presence of two-dimensional adverse selection; and (3) the presence of moral hazard and non-separable preferences (Bennardo & Chiappori 2003).

each insured can purchase only one policy. Finally, the insured purchases his or her policy in period 0 and pays a level premium that amounts to γ units of output each period. None of the insured will default on his or her premiums.

6.2 Pure Moral Hazard Model

The pure moral hazard model assumes that ex ante none of the insured has suicide intention. However, ex post some of the insured suffer permanent shocks that lead them to consider committing suicide. Committing suicide is a kind of hidden action that cannot be observed by insurance companies. Thus, moral hazard arises because suicide exclusion would affect the timing when the insured commits suicide. The insured has control over the timing of his or her suicide so he or she would defer committing suicide until it becomes a covered risk.

To formalize this model, I assume that household heads are identical in period 0, and none of them have suicide intention at that time. However, although they are endowed with one unit of labor individually, production uncertainty makes them lucky or unlucky beginning in period 1. In period 1, lucky household heads, whose proportion is (1-q), will produce Y units of output for each member in the household. However, unlucky household heads, whose proportion is q, will suffer permanent shocks and consequently produce αY units of output for each member in the household. The value of α varies among unlucky household heads and is uniformly distributed between $\underline{\alpha}$ and $\overline{\alpha}$ ($1>\overline{\alpha}>\underline{\alpha}>0$). Furthermore, for each unlucky household head, his or her value of α remains the same in each period after the shock. In period 2, unlucky household heads remain unlucky, but a proportion q of the lucky household heads will become unlucky. In each period after period 2, unlucky household heads remain unlucky, but a proportion q of the lucky household heads will become unlucky.

Unlucky household heads suffer low output caused by permanent shocks. Therefore, they may consider committing suicide so that their death benefits could be secured earlier even though this action will incur a utility loss of $u(1+a)$, where a is the cost of committing suicide[27]. Additionally, for the sake of tractability, I make the following two assumptions. First, even the unluckiest household heads can afford life insurance. That is, $\underline{\alpha}Y-\gamma \geq 0$. Secondly, utility gain can be secured if the insured commits suicide and death benefits are received one period later. That is, $-u(1+a)+\beta u(B) > 0$.

[27] For the insured who does not commit suicide, $a=0$ and $u(1)$ is normalized to 0.

On the supply side, insurance companies choose the premium that maximizes their expected profits, and Bertrand competition drives the expected profits to be zero in equilibrium. That is,

$$(1)\ \gamma + \gamma\left\{ \sum_{t=1}^{t=T-1} \frac{\left[1-\left(\sum_{s=1}^{s=t} p_s\right)\right]}{(1+R)^{t+1}} \right\} = B\left[\sum_{t=1}^{t=T-1} \frac{p_t}{(1+R)^{t+1}}\right] + B\frac{\left[1-\sum_{s=1}^{s=T-1} p_s\right]}{(1+R)^{T+1}}$$

where equation (1) represents the zero profit constraint; p_s represents the suicide rate of the insured in period *s*; and *R* represents the interest rate.

Since insurance markets are competitive, household heads will be better off purchasing life insurance. Consequently, on the demand side, all of the household heads will purchase life insurance. By taking the premium value as given, household heads choose their optimal actions (committing suicide or staying alive) that maximize their expected utility. For the sake of reality, I assume that lucky household heads will never commit suicide even when suicide becomes a covered risk. That is,

(2) $u(Y-\gamma)+\beta\ E_t\ [U(\alpha,\gamma)] \geq u(Y-\gamma) - u(1+a)+\beta u(B),\ 2\leq t\leq T-1.$

where $E_t\ [U(\alpha,\gamma)]$ represents the expected utility of the lucky insured in period t.

I solve the optimal action of the unlucky insured by backward induction. In period T, since everyone will die, and committing suicide incurs a utility loss, none of the insured will choose to commit suicide. In period (T-1), unlucky household heads will choose to commit suicide if their utility gain from staying alive is outweighed by their utility gain from committing suicide. That is,

$$\beta u(\alpha Y-\gamma)+\beta^2 u(B) \leq - u(1+a)+\beta u(B)$$

Let the equality hold when the value of α equals α_{T-1}. Thus,

$\alpha_{T-1}=\frac{\gamma}{Y}+\frac{1}{Y}*u^{-1}[\frac{-1}{\beta}u(1+a)+(1-\beta)u(B)]$, and unlucky household heads whose value of α is less than α_{T-1} will commit suicide in period (T-1).

In period (T-2), it is straightforward to show that unlucky household heads whose value of α is less than α_{T-1} will commit suicide. For unlucky household heads whose value of α is greater than α_{T-1}, they will commit suicide if their utility gain from living for two more periods is outweighed by their utility gain from committing suicide now. That is,

$\beta\, u(\alpha\, Y\text{-}\gamma) + \beta^{2} u(\alpha\, Y\text{-}\gamma) + \beta^{3} u(B) \leq -u(1+a) + \beta\, u(B)$

Let the equality hold when the value of α equals α_{T-2}. Thus,

$\alpha_{T-2} = \frac{\gamma}{Y} + \frac{1}{Y} * u^{-1}[\frac{-1}{(\beta+\beta^{2})} u(1+a) + (1-\beta)u(B)]$, and unlucky household heads whose value of α is less than α_{T-2} will commit suicide in period (T-2).

It is straightforward to show that the value of α_{T-2} is greater than that of α_{T-1}. This suggests that the cutoff value of α, which makes unlucky agents indifferent between committing suicide and staying alive, in period (T-2) is greater than that in period (T-1). The reason that α_{T-2} is greater than α_{T-1} can be explained as follows: in period (T-1), the unlucky household heads whose value of α equals α_{T-1} are indifferent between committing suicide and staying alive for one more period (in period T). This suggests that committing suicide is superior to staying alive for two more periods (in period T-1 and T) because the utility gain from earning the low output for an additional period is outweighed by the utility loss from receiving death benefits one period later. That is, $\beta\, u(\alpha_{T-1} Y\text{-}\gamma) < (\beta-\beta^{2})\, u(B)$. Consequently, the unlucky household heads whose value of α equals α_{T-1} will strictly prefer to commit suicide in period (T-2). Thus, to make household heads indifferent in period (T-2), output has to be increased. This means that the cutoff value of α has to be increased from α_{T-1} to α_{T-2}.

By the same method, I derive the cutoff value of α in period (T-t) to be

$\alpha_{T-t} = \frac{\gamma}{Y} + \frac{1}{Y} * u^{-1}[\frac{-1}{(\sum_{s=1}^{s=t} \beta^{s})} u(1+a) + (1-\beta)u(B)]$, where $1 \geq t \geq$ (T-2).

It is straightforward to show that between period 2 and (T-1), the cutoff value of α declines over periods. Furthermore, in period 1, since suicide is not a covered risk, none of the unlucky insured will choose to commit suicide. Let $F(\alpha_{t})$ represent the probability that the value of α is less than α_{t}, the cutoff value of α in period t. This enables me to calculate the suicide rate, which is measured by the proportion of household heads committing suicide, by multiplying $F(\alpha_{t})$ by its correspondent proportion of unlucky household heads. The results, which are summarized in the third row of the table below, show that the suicide rate increases from period 1 to period 2, declines from period 2 on, and reaches 0 in period T.

Period	1	2	3	t	T
Cutoff Value of α	N/A	α_2	α_3	α_t	N/A
Suicide Rate	0	$F(\alpha_2)$* [q+(1-q)*q]	$F(\alpha_3)$* $[(1-q)^2*q]$	$F(\alpha_t)$* $[(1-q)^{t-1}*q]$	0

Finally, I substitute the values of suicide rates into equation (1), the zero profit constraint, to derive the equilibrium value of premiums. Since it is too complicated to derive a general formula of the value, I employ a numeric example to show the premium value. I assume a logarithmic utility function and illustrate the numeric values of the relevant parameters in the table below[28].

Parameter	Value	Parameter	Value
Number of Periods (T)	25	Death Benefits (B)	140,000
Proportion of Unlucky Household Heads (q)	10^{-3}	Utility Loss of Committing Suicide (a)	500
Output of Lucky Household Heads (Y)	24,550	Intertemporal Discount Factor (β)	0.85
Measurement of Unluckiness (α)	Uniformly Distributed between 0.015 and 0.025	Interest Rate (R)	0.176

Using the preceding parameter values, I calculate the equilibrium value of premiums to be 368. Furthermore, I show the suicide rate of each period in Figure 3. This figure shows that the suicide rate increases significantly in period 2, declines by half in period 3, and then gradually declines to 0 in the 25th period.

6.3 Model that Combines Adverse Selection and Moral Hazard

The model that combines adverse selection and moral hazard assumes that ex ante some of the insured has suicide intention, and they choose to commit suicide after it becomes a covered risk. When suicide intention is a kind of public information, insurance companies can discriminate these heterogeneous agents by offering two types of policies. On the one hand, companies can offer policies with low premiums to the agents who do not have suicide intention.

[28] From 1990 to 1995 in the United States, the amount of life insurance coverage per insured household averaged $140,917, and the disposable personal income per household averaged $49,100 (American Council of Life Insurance 1996).

On the other hand, companies can offer policies with extremely high premiums to agents who have suicide intention. However, when suicide intention is a kind of hidden information, companies will incur losses if they offer the preceding two types of policies since both types of agents will choose policies with low premiums.

Rothschild and Stiglitz (1976) propose a screening model to explore the problems that may arise when hidden information is present in competitive insurance markets. First, they argue that a pooling equilibrium, in which identical policies are purchased by heterogeneous agents, cannot exist. Secondly, they argue that if an equilibrium exist, it must be a separating equilibrium in which companies offer a menu of policies that make agents reveal their own types and choose their own policies.

The model of Rothschild and Stiglitz (1976) assumes no administrative costs in marketing a menu of policies. However, after taking into account administrative costs and state regulations limiting the length of suicide exclusion, insurance companies actually provide identical policies on suicide exclusion even though the insured may have heterogeneous suicide intention ex ante. This suggests that in practice, a pooling equilibrium may exist. Since my model serves to characterize the incentive effect of the suicide exclusion, I assume that insurance companies offer identical policies and a pooling equilibrium exists.

In this model, the basic setup of households and insurance markets remains the same as those outlined in Section 3.1. However, I assume that two types of household heads exist ex ante. Type 1 household heads, whose proportion is (q), have suicide intention in period 0. This intention is a kind of hidden information, only known to the insured. Type 2 household heads, whose proportion is (1-q), do not have suicide intention at all. These two types of household heads differ in suicide intention only. Suicide intention reduces the cost to commit suicide. Thus, I assume that $a_2 >> a_1 > 0$, where a_1 is the cost of committing suicide for type 1 household heads, while a_2 is the cost of committing suicide for type 2 household heads.

The supply side of the insurance market is the same as those outlined in Section 3.2. Insurance companies choose the premium that maximizes their profits, and Bertrand competition drives the profits to be zero in equilibrium [Equation (1) in Section 3.2].

In addition, the demand side of the insurance market resembles those outlined in Section 3.2. All of the household heads will purchase life insurance and choose their optimal actions (committing suicide or staying alive) that maximize their utility. For the sake of reality, I assume that type 2 household heads will never commit suicide even when suicide becomes a covered risk. That is, $u(Y-\gamma)+\sum_{s=t+1}^{s=T}\beta^{s-t}u(Y-\gamma) \geq u(Y-\gamma) - u(1+a_2)+\beta u(B)$, $2 \leq t \leq$ T-1. Additionally, I assume that a pooling equilibrium exists. I solve the optimal action of type 1 household heads by backward induction. In period T, since everyone will die, and committing suicide incurs a utility loss, none of the household heads will choose to commit suicide. In period (T-1), type 1 household heads will choose to commit suicide if their utility gain from staying alive is outweighed by their utility gain from committing suicide. That is,

$$\beta u(Y-\gamma)+\beta^2 u(B) \leq - u(1+a_1)+\beta u(B)$$

Let the equality hold when the value of γ equals γ_{T-1}. Thus,

γ_{T-1} =Y- $u^{-1}[\frac{-1}{\beta}u(1+a_1)+(1-\beta)u(B)]$, and any premium values greater than γ_{T-1} will induce type 1 household heads to commit suicide in period (T-1).

In period (T-2), it is straightforward to show that any premium values greater than γ_{T-1} will induce type 1 household heads to commit suicide. For premium values less than γ_{T-1}, they will induce type 1 household heads to commit suicide if the heads' utility gain from living for two more periods is outweighed by their utility gain from committing suicide now. That is,

$$\beta u(Y-\gamma)+\beta^2 u(Y-\gamma)+\beta^3 u(B) \leq - u(1+a_1)+\beta u(B)$$

Let the equality hold when the value of γ equals γ_{T-2}. Thus,

γ_{T-2} =Y- $u^{-1}[\frac{-1}{(\beta+\beta^2)}u(1+a_1)+(1-\beta)u(B)]$, and any premium values greater than γ_{T-2} will induce type 1 household heads to commit suicide in period (T-2).

It is straightforward to show that the value of γ_{T-2} is less than that of γ_{T-1}. This suggests that the cutoff value of γ, which makes type 1 household heads indifferent between committing suicide and staying alive, in period (T-2) is less than that in period (T-1). The reason that γ_{T-2} is less than γ_{T-1} can be explained as follows: when premiums equal γ_{T-1} in period (T-1), type 1 household heads are indifferent between committing suicide and staying alive for one more period (in period T). This suggests that committing suicide is superior to staying alive for two more periods (in period T-1 and T) because the utility gain from consumption in an additional period is outweighed by the utility loss from receiving death benefits one period later. That is,

$\beta\, u(Y- \gamma_{T-1}) < (\beta - \beta^2)u(B)$. Consequently, if premiums equal γ_{T-1} in period (T-2), type 1 household heads will strictly prefer to commit suicide. Thus, to make type 1 household heads indifferent in period (T-2), consumption has to be increased. This means that the cutoff value of γ has to be reduced from γ_{T-1} to γ_{T-2}.

By the same method, I derive the cutoff value of γ in period (T-t) to be

$$\gamma_{T-t} = Y - u^{-1}\left[\frac{-1}{\left(\sum_{s=1}^{s=t} \beta^s\right)} u(1+a_1) + (1-\beta)u(B)\right], \text{ where } 1 \geq t \geq (T-2).$$

It is straightforward to show that between period 2 and (T-1), the cutoff value of γ increases over periods. Furthermore, in period 1, since suicide is not a covered risk, and committing suicide incurs a utility cost, none of the type 1 household heads will choose to commit suicide.

The preceding analysis suggests the relationship between premiums and the suicide behavior of the type 1 household heads. If the premium is set so low that its value is less than the cutoff value in period 2 (γ_2), none of the type 1 household heads will commit suicide throughout the periods. However, if the premium is set above γ_2, all of the type 1 household heads will commit suicide in period 2. In summary, the analysis predicts that depending on premium values, the suicide rate should either be zero in all periods or resemble a spike with a positive rate in period 2 but zero rates in the other periods.

I assume a logarithmic utility function and employ a numeric example, whose parameter values are in the table below, to illustrate the preceding prediction.

Parameter	Value	Parameter	Value
Number of Periods (T)	25	Death Benefits (B)	160,000
Proportion of Type 1 Household Heads (q)	10^{-5}	Intertemporal Discount Factor (β)	0.85
Output of Household Heads (Y)	24,000	Interest Rate (R)	0.176
Type 1 Household Heads' Utility Loss of Committing Suicide	500	Type 2 Household Heads' Utility Loss of Committing Suicide	10,000

Using the preceding parameter values, I calculate the equilibrium premium value to be 420 and the cutoff premium value in period 2 (γ_2) to be 90. Since the equilibrium premium value exceeds the cutoff premium value in period 2, this suggests that all of the type 1 household heads will commit suicide in period 2. The suicide rate is 10 suicides per 1,000,000 household heads in period 2 and zero in the other periods.

6.4 Testable Implications

Both the pure moral hazard model and the model that combines adverse selection and moral hazard predict that the suicide rate should be positive in period 2. These predictions can be used to show presence of asymmetric information that suggests that the suicide rate should be positive in period 2 (the period right after the expiration of the suicide exclusion). Moreover, these two models have different predictions regarding the suicide rate in later periods. On the one hand, the model that combines adverse selection and moral hazard predicts that the suicide rate should be zero after period 2. On the other hand, the pure moral hazard model predicts that the suicide rate should remain positive after period 2. These two different predictions can be used to disentangle the two models under some conditions. If the observed suicide rate becomes zero after period 2, this suggests that the pure moral hazard model should be rejected. If the observed suicide rate declines but remains positive after period 2, this suggests that neither of the two models should be rejected since this pattern could be caused by both models or by the pure moral hazard model alone.

7. Conclusion

This paper investigates the incentive effect of a two-year suicide exclusion on individual life insurance policies by employing aggregate mortality data provided by the Society of Actuaries. Although the nature of the acquired data limits the scope of this paper, the analyses yield two findings uniformly evident across genders and age groups; these findings suggest that the suicide rate quadruples after the exclusion period and that suicides during the exclusion period are often disguised as accidental deaths. Therefore, the findings provide evidence of asymmetric information and imply that economic considerations affect the timing and method of committing suicide.

The preceding results can help insurance companies price mortality risks resulting from the implementation of the suicide exclusion. Additionally, by estimating the loss to the beneficiaries of the insureds committing suicide, this paper suggests an annual loss of $30 billion (in 2001 dollars) to the survivors of the people dying by suicide in the United States[29]. Although this loss value is a crude measure, it sheds light on the benefits yielded by the implementation of the *National Strategy for Suicide Prevention*, the plan that guides suicide prevention efforts in the United States.

[29] The estimation method of this loss value is detailed in the data appendix. Similarly, Goldsmith, Pellmar, Kleinman, and Bunney (2002) suggest that suicides in the United States in 1998 incurred productivity losses of 11.8 billion (in 1998 dollars). Miller, Covington, and Jensen (1999) suggest that suicides in the United States in 1995 incurred medical expenses of $3.7 billion, work-related losses of $27.4 billion, and quality of life costs of $80.2 billion.

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Data Appendix

While 20 companies contributed data to the 1990-1995 SOA mortality studies, only 9 companies submitted cause-of-death information. As mentioned in Section 3, these data tend to represent larger companies and companies with lower mortality rates. I discuss below the aforementioned aspects of the data and the issue of data accuracies. I also outline three estimation methods that are used to calculate the upper bound value of the selection effects (Section 4), the annual loss to the survivors of the people dying by suicide (Section 6), and the numbers of policies terminated by suicide and by accidental deaths (Table 3 and 5).

First, to save administrative costs in collecting data, the SOA chooses to collect data from large companies. In terms of the coverage of life insurance policies in force, these 9 companies submitting cause-of-death information ranked among the leading 200 companies of the 1,034 companies in the United States in 1996. Their market share amounted to 11% (A.M. Best Company, 1997). Secondly, companies vary in the coding practice of cause-of-death information. While others do not, some companies invest a certain amount of resources in information technology so that they can keep a good internal cause-of-death record. This record puts companies in a better position to remedy their deficiencies in the underwriting process and therefore may have an impact on lowering mortality rates. In the 1990-1995 SOA studies, 9 companies could supply cause-of-death information, while the other 11 could not. This suggests that these 9 companies might have had lower mortality rates when other factors affecting mortality rates, such as the segment of the targeted market, were controlled.

Additionally, the cause-of-death information is the key variable in this paper, so it is important to consider its accuracy. The 1990-1995 SOA mortality studies report death rates by cause of death in eleven broad categories: four external causes (suicide, homicide, motor vehicle accidents, and all other accidents); six natural causes (diseases of heart and circulation systems, cerebro-vascular diseases, malignant neoplasms, diabetes mellitus, pneumonia and influenza, and AIDS); one category of all other causes. Admittedly, the reported information may not have accurately mirrored the true underlying cause of death of the insured. This inaccuracy could have been caused by several factors, such as the variation in companies' coding practices and the variation in state regulations in completing the cause of death on the death certificate. However, any inaccuracies seem to have affected the death rates for causes contributing to natural deaths, but not to have affected the death rate for suicide. This is because of the standard procedure for reporting cause-of-death information, which classifies deaths according to 6 categories in the section of the death certificate called “manner of death”: natural deaths, accidents, suicide, homicide, pending investigation, and “could not be determined”. When natural deaths occur, the certifying physician should provide three kinds of cause-of-death information on the death certificate including immediate cause, underlying cause, and other significant conditions. According to Jack Luff, the actuary at the SOA in charge of the 1990-1995 mortality studies, companies are requested to report only one cause, the underlying cause, to the studies for each policy terminated by death. However, it is not clear whether companies are always reporting the underlying cause. If some companies report the immediate cause, and the underlying and immediate causes fall into different categories, this will create measurement errors in death rates for causes contributing to natural deaths. For example, an insured individual dies from complications of malignant neoplasms, such as pneumonia. In this case, malignant neoplasms are the underlying cause, while pneumonia is the immediate cause. If some companies mistakenly report pneumonia as the underlying cause, this will cause measurement errors because the death rate due to malignant neoplasms would be underestimated, while the death rate due to pneumonia would be overestimated.

Moreover, some states impose strict rules on reporting cause-of-death information, while others do not. This variation may also cause inaccuracies in death rates for causes contributing to natural deaths. For example, during the period from the late 1980s to the early 1990s, some states required that if AIDS was the underlying cause, it should be accurately reported. This requirement was imposed because the stigma attached to AIDS prevented certifying physicians from honestly reporting AIDS, but instead simply reporting its complications, such as pneumonia. Thus measurement errors were more likely to arise in the data collected from the states that did not impose this requirement.

In summary, the preceding two types of variation affect the death rates for causes contributing to natural deaths because they transfer deaths across cause-of-death categories. However, these two types of variation are not relevant to the death rate for suicide because when deaths suspected to be suicide occur, they must be referred to the medical examiner or coroner who will investigate the deaths and decide whether or not they are due to suicide. If the deaths are due to suicide, the medical examiner or coroner will mark them as suicide in the “manner of death" section of the death certificate. Since the marked category in the "manner of death" section matches the cause-of-death category that the staff in the claims department of insurance companies should report, this enables the staff to simply code suicide directly from the "manner of death" section of the death certificate. As a result, this procedure reduces the factors that may cause measurement errors in the death rate for suicide.

In Section 4, the calculation of the upper bound value of the selection effects suggests that the male natural death rate during the 3rd-4th policy years is at most 1.7 times higher than that during the first two policy years. This calculation is performed by using three categories of data: (1) natural death rates of male insureds during the 1st-2nd policy years and during the 7th-8th policy years; (2) the distribution of exposure by issue age and by policy duration; (3) natural death rate of male adults in the general population. The second category of data is used to calculate the proportions of healthy and unhealthy insureds, and the third category of data is used to control age effects.

Additionally, Section 6 estimates the annual loss to the survivors of the people dying by suicide. This estimation assumes a participation constraint in which the insureds commit suicide when their valuation of their own lives and death benefit payments exceeds their valuation of the loss to their beneficiaries. If a "zero value of life" is assumed, the value of death benefit payments serves as the upper bound value of the loss to the beneficiaries. Furthermore, if survivors' loss does not depend on the availability of death benefit payments, the value of death benefit payments can be used to calculate the upper bound value of the loss to the survivors of the people dying by suicide. Therefore, when the participation constraint is binding, the annual loss to the survivors is estimated to be $29 billion (in 2001 dollars) due to two pieces of information: (1) 30,900 people died by suicide each year in 1990-1995; (2) the statistics acquired from litigation reports suggest that the amount of insurance coverage owned by an insured person committing suicide averages $933,000 (in 2001 dollars)[30].

Finally, Table 3 reports (in parentheses) the estimated numbers of policies terminated by

[30] The insured often owns multiple life insurance policies, and the SOA data do not include the information on the overall amount of coverage that an insured person owns. Therefore, I turn to litigation reports to seek the overall amount of coverage owned by an insured person committing suicide. These reports include Nolan (1988), Mealey Publications and Conferences (1998), Maniloff (2002), Wallin v. Ins. Co. of North America and Federal Guaranty Life Ins. Co. (1980), United Fence Co. v. Great-West Life Assurance Co. (1986), Lange v. Penn Mutual Life Insurance Co. (1988), and Sonderegger v. United Investors Life Ins. Co. (1992).

suicide. These numbers are estimated from four categories of data: (1) the total number of policies terminated by suicide (7,130 policies); (2) the distribution of exposure by gender, by issue age, and by duration; (3) the distribution of policy size by duration[31]; (4) the by-amount suicide rates. Furthermore, I calculate the ratio between the by-amount suicide rate and the by-amount accidental death rate. I use this ratio, the preceding first three categories of data, and the by-amount accidental death rates to estimate the number of policies terminated by accidental deaths. These estimated numbers are reported in the parentheses of Table 5.

[31] *Life Insurance Fact Book* and *Statistical Abstracts of the United States* report the amount of coverage and the number of ordinary life insurance policies purchased each year in the United States. These two kinds of information are used to calculate the distribution of policy size by duration.

Table 1: Summary Statistics of the Cause-of-Death Data in the 1990-1995 SOA Mortality Studies[1]

1. Number of Policy Records[2]: 48.89 million

2. Number of Policies Terminated By Death

Category of Death	Suicide	All Causes
No. of Policies	7,130	381,212
%	1.87	100

3. Amount of Insurance Coverage Exposed to Loss (Exposure): $2.7 trillion

4. Distribution of Exposure

4.1 By Gender: Male: 72%
Female: 28%

4.2 By Issue Age

Issue Age	0-19	20-29	30-39
%	9	18.4	39

Issue Age	40-49	50-59	60 & up
%	23	8.2	3

4.3 By Policy Duration

Policy Years	1st-2nd	3rd-4th	5th-6th
%	30	21.5	15.1

Policy Years	7th-8th	9th & up
%	10	22.8

1. Data Sources: The information on the number of policy records and the number of policies terminated by death comes from unpublished data provided by the Society of Actuaries. The statistics on the amount and distribution of exposure are estimated from the reports and tables of the 1990-1995 studies on mortality under standard individually underwritten life insurance. These reports and tables provide statistics on the amount and distribution of exposure contributed by all of the twenty companies and on the percentage of contribution by each individual company. I use these statistics to estimate the amount and distribution of the nine companies contributing the cause-of-death information by assuming that these nine companies have the same distribution of exposure as the overall twenty companies.
2. The term of policy records takes into account the number of years that a policy is in force during the 1990-1995 period. For example, the number of policy records is 2 for a policy in force for two years.

Table 2: Suicide Rates of the General Population and the Insured

(1) General Population				(2) Insured Whose Policies are within the First Ten Policy Years				
Age of Death	Suicide Rate Male	Suicide Rate Female	Gender Ratio	Age of Issue	Age of Death	Suicide Rate Male	Suicide Rate Female	Gender Ratio
15~24	22.80	3.87	5.90	10-19	15~24	12.25	1.81	6.78
25~34	25.60	5.41	4.73	20-29	25~34	5.36	2.16	2.48
35~44	24.30	6.46	3.76	30-39	35~44	5.94	2.21	2.69
45~54	22.54	7.05	3.19	40-49	45~54	9.24	3.13	2.96
55~64	24.49	6.23	3.93	50-59	55~64	9.69	2.81	3.44
65 & up	39.46	5.81	6.80	60-79	65~84	11.09	3.22	3.45
All Ages	20.19	4.65	4.34	All Ages	All Ages	6.99	2.09	3.34

1. General population and the insured have different units of suicide rates. For the general population, the suicide rate is measured by the number of people who die by suicide per 100,000 people. For the insured, the suicide rate is measured by the amount of death benefits paid out to suicides per $100,000 insurance coverage exposed to loss. Consequently, the suicide rate of the insured can be viewed as the weighted suicide rate of the general population. The weight is the amount of coverage.

2. Gender ratio represents the ratio between the male suicide rate and the female suicide rate.

3. Derived from the 1990-1995 SOA mortality studies, the suicide rate of the insureds represents the weighted average suicide rate of the insureds whose policies are within the first ten policy years; the weight is the distribution of exposure by policy duration. Similarly, the suicide rate of the general population represents the yearly average of the suicide rate in 1990-1995.

4. Data sources of suicide rates of the general population: the data in 1990-1993 come from the Statistical Abstracts of the United States (1993-1996), the 1994 data come from the U.S. Monthly Vital Statistics Report (1995), and the 1995 data come from the website of the National Center for Health Statistics, http://www.cdc.gov/nchs/datawh/statab/unpubd/mortabs/gmwkiii.htm.

5. From 1990 to 1995 in the United States, the annual average of the resident population amounted to 256 million, and the annual average of the number of ordinary life insurance policies in force amounted to 141 million.

Table 3: Suicide Rates of the Insured

Issue Age	Death Rate by Policy Years				
	1st-2nd	3rd-4th	5th-6th	7th-8th	All Years (1st-8th)
(1) Male					
20-29	1.27	6.43	9.45	5.31	4.88
	(17)	(68)	(80)	(36)	(201)
30-39	1.59	7.21	8.62	7.99	5.40
	(46)	(162)	(154)	(115)	(477)
40-49	3.46	11.25	11.27	11.35	8.23
	(58)	(147)	(117)	(95)	(417)
50-59	1.73	11.23	12.50	8.64	7.44
	(11)	(53)	(47)	(26)	(137)
All Ages	2.02	8.46	9.83	8.36	6.21
(20-59)	(132)	(430)	(398)	(272)	(1232)
(2) Female					
20-29	0.73	3.15	2.20	2.00	1.87
	(4)	(13)	(7)	(2)	(26)
30-39	0.36	1.93	4.07	2.45	1.81
	(4)	(17)	(28)	(14)	(63)
40-49	0.73	2.21	6.28	5.58	2.88
	(5)	(11)	(25)	(18)	(59)
50-59	0.12	2.09	8.13	0.18	2.28
	(1)	(4)	(12)	(1)	(18)
All Ages	0.51	2.27	4.62	2.97	2.15
(20-59)	(14)	(45)	(72)	(35)	(166)

1. The death rate is measured by the amount of death benefits paid out per $100,000 insurance coverage exposed to loss. This measure is discussed in Section 4.1.

2. Numbers in parentheses are the estimated numbers of policies terminated by suicide. The estimation method is outlined in the data appendix.

Table 4: Estimates of the Incentive Effect of the Suicide Exclusion
Dependent Variable: ln (Suicide Rates)

Explanatory Variables	(1) OLS	(2) WLS	(3) OLS	(4) WLS	(5) OLS	(6) WLS
	1. Both Genders		2. The Male		3. The Female	
Duration Dummy	1.67	1.45	1.39	1.32	1.92	1.79
(=1 if Duration > 2 Years)	(0.32)	(0.16)	(0.16)	(0.13)	(0.62)	(0.38)
Female Dummy	-1.40	-1.25				
	(0.26)	(0.16)				
ln (Natural Death Rates)	0.01	0.16	0.17	0.24	-0.11	-0.01
	(0.11)	(0.08)	(0.06)	(0.06)	(0.21)	(0.16)
Intercept	0.51	0.16	0.11	-0.02	-0.62	-0.76
	(0.44)	(0.24)	(0.21)	(0.20)	(0.77)	(0.49)
R^2	0.69	0.87	0.91	0.93	0.43	0.68
Sample Size	32	32	16	16	16	16

1. The observations come from Table 3 that includes the insureds whose issue ages are between 20 and 59 and whose policies are within the first eight years.
2. In Column (2), (4), and (6), the distribution of exposure is used as the weight to perform the weighted least squares estimation.
3. Numbers in parentheses are standard errors.

Table 5: Accidental Death Rates of the Insured

Issue Age	Death Rate by Policy Years				
	1st-2nd	3rd-4th	5th-6th	7th-8th	All Years (1st-8th)
(1) Male					
20-29	16.23	16.01	13.91	15.05	15.55
	(325)	(247)	(164)	(149)	(885)
30-39	12.56	10.87	16.06	12.50	12.77
	(513)	(343)	(399)	(245)	(1500)
40-49	19.71	12.69	19.99	12.49	16.86
	(456)	(229)	(285)	(146)	(1116)
50-59	20.74	13.31	25.83	36.28	21.69
	(166)	(91)	(134)	(168)	(559)
All Ages	15.93	12.64	17.51	15.17	15.23
(20-59)	(1460)	(910)	(982)	(708)	(4060)
(2) Female					
20-29	5.05	4.03	5.15	4.24	4.68
	(42)	(25)	(25)	(17)	(109)
30-39	5.87	4.11	7.13	4.12	5.40
	(99)	(51)	(71)	(33)	(254)
40-49	9.02	7.45	6.81	3.17	7.38
	(87)	(55)	(39)	(16)	(197)
50-59	7.05	7.15	13.27	2.25	7.70
	(24)	(21)	(28)	(4)	(77)
All Ages	6.62	5.23	7.18	3.73	5.97
(20-59)	(252)	(152)	(163)	(70)	(637)

1. The death rate is measured by the amount of death benefits paid out per $100,000 insurance coverage exposed to loss. This measure is discussed in Section 4.1.

2. Numbers in parentheses are the estimated numbers of policies terminated by accidental deaths. The estimation method is outlined in the data appendix.

Table 6: Rates of Overall Suicides Resulting from the Implementation of the Suicide Exclusion

	Difference by Cause of Death			
Issue Age	Suicide	Accidental Deaths	Overall Suicides	% of Disguised Suicides
(1) Male				
20-29	5.16	-0.22	5.38	4%
30-39	5.62	-1.69	7.31	23%
40-49	7.79	-7.02	14.81	47%
50-59	9.5	-7.43	16.93	44%
All Ages (20-59)	6.44	-3.27	9.71	34%
(1) Female				
20-29	2.42	-1.02	3.44	30%
30-39	1.57	-1.76	3.33	53%
40-49	1.48	-1.57	3.05	51%
50-59	1.97	0.1	1.87	N/A
All Ages (20-59)	1.76	-1.4	3.16	44%

1. Difference in death rates = (the death rate during the 3rd-4th policy years)-(the death rate during the 1st-2n policy years)

2. The rate of overall suicides is calculated by taking the difference between the difference in suicide rates and the difference in accidental death rates.

3. The percentage of disguised suicide is calculated by dividing the decline in accidental death rates between the 1st-2nd policy years and the 3rd-4th policy years by the rate of overall suicides.

4. The death rate is measured by the amount of death benefits paid out per $100,000 insurance coverage exposed to loss. This measure is discussed in Section 4.1.

Figure 1: Suicide Rates of the Insured

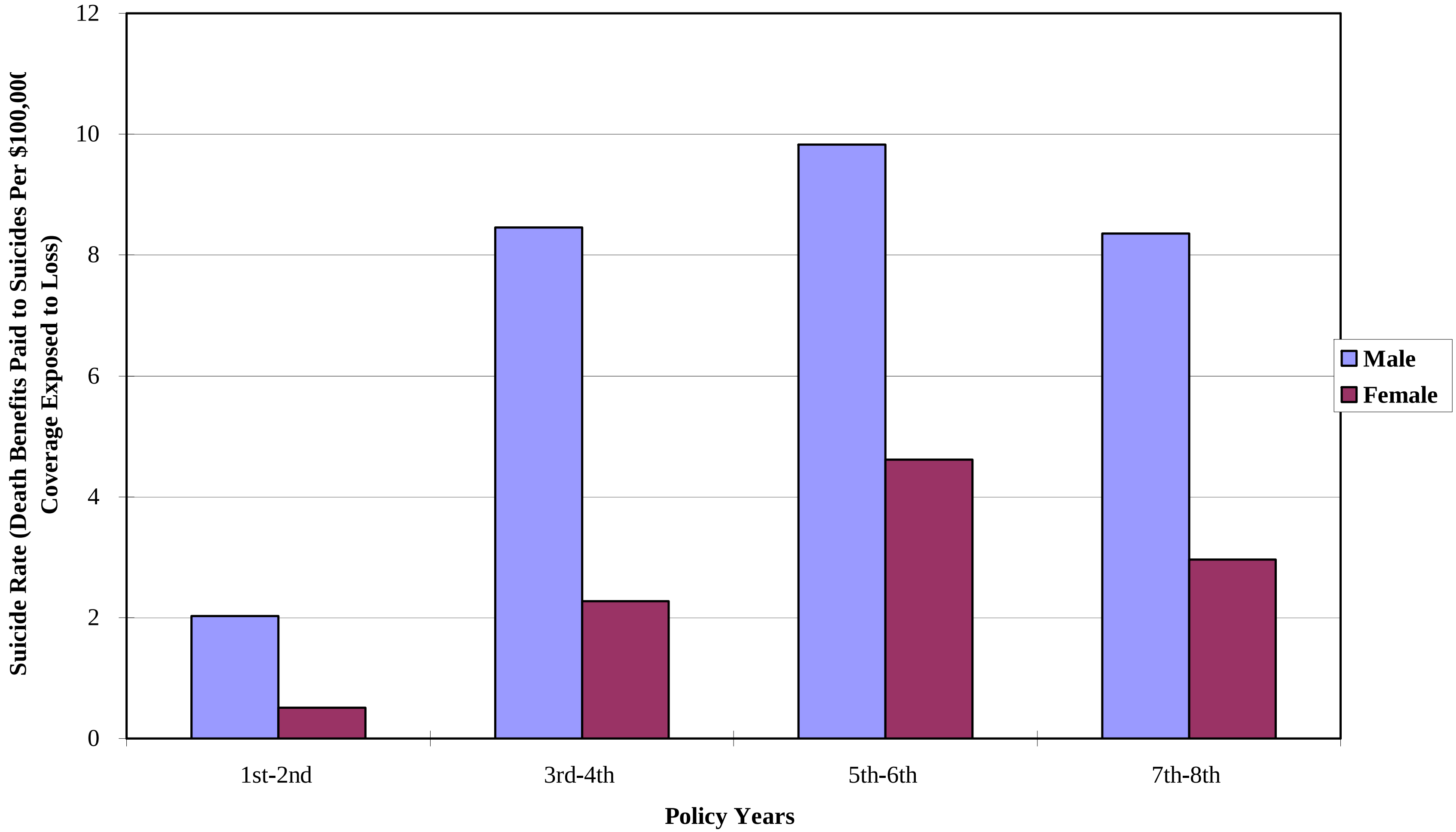

Figure 2: Natural Death Rates of the Insured

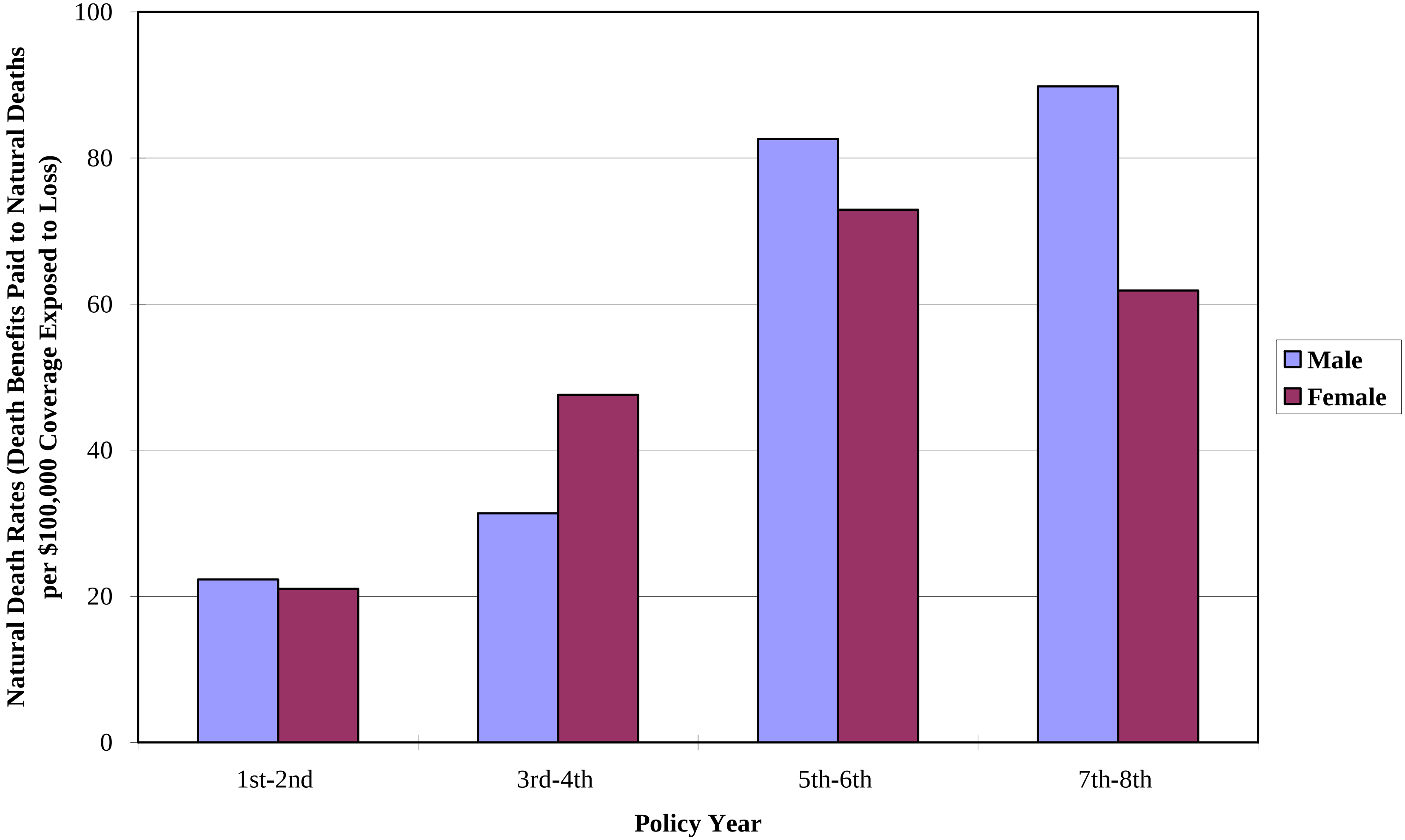

Figure 3: Suicide Rates Derivded from the Pure Moral Hazard Model

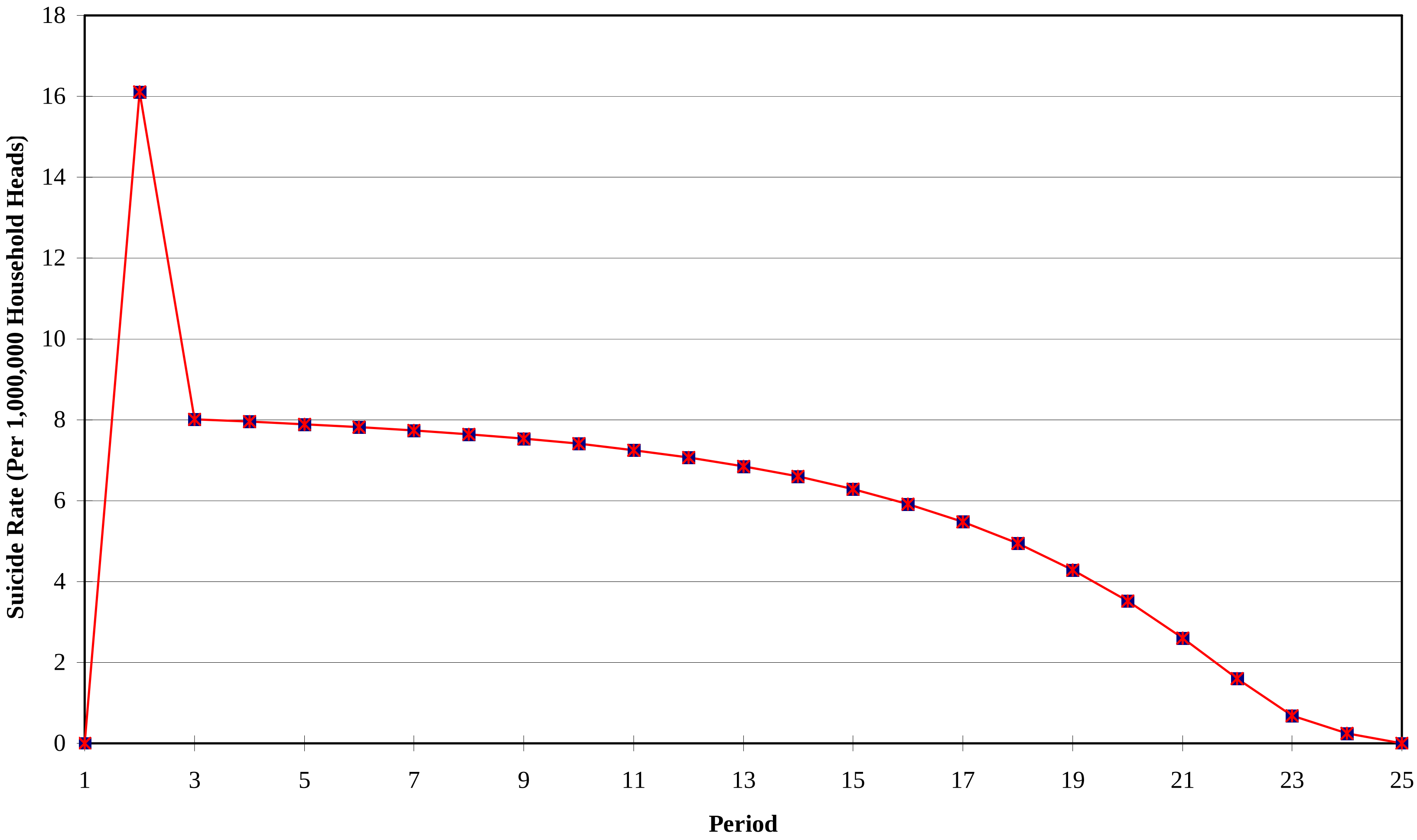

Figure 4: Suicide Rates Derived from the Model that Combines Adverse Selection and Moral Hazard