On the Demand Effects of Rate Regulation – Evidence from a Natural Experiment

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Abstract

We analyze the influence of rate regulation on insurance demand in an annuity setting. With a unique dataset containing a natural experiment due to German federal regulation and the E.U. Gender Directive we study the impact of unisex tariffs on contract choices in variable annuity products. Our data contains two different choice variables with antithetic predictions for men and women, meaning that women should increase their demand in one choice and decrease it in the other, while men should exhibit opposite behavior. We find with regard to both choices that both men and women have lower demand for guarantees within the annuity in unisex contracts than without rate regulation. This behavior contradicts economic intuition. We hypothesize that the effect could instead be explained by the public perception of unisex tariffs.

Keywords: Adverse selection, law and economics, rate regulation, unisex tariffs

JEL-Classification: D14, D81, D82, G22, K20, L51

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1 Introduction

The impact of rate regulation in insurance markets has been discussed intensively over the last decades. Specifically the issue of gender neutral tariffs, i.e. the use of unisex tariffs, has been a continuous issue of policy debates. Such legislation induces non-adequate pricing of the contracts and thus has the potential of causing adverse selection, which causes inefficiencies from a welfare perspective (Crocker and Snow, 1986; Rea, 1987; Rothschild, 2011). This argument, however, is only viable under the assumption that consumers adjust their demand when the pricing of an insurance contract is changed. Whether this assumption is valid is the topic of this study.

Rate regulation banning gender based tariffs exists both in the United States and the European Union due to fairness considerations.[1] In the U.S., two Supreme Court decisions in 1978 and 1983 prohibit the use of separate mortality tables for men and women in pension benefit calculations due to the legal definition of discrimination in the Civil Rights Act of 1964 (McCarthy and Turner, 1993). In the E.U., gender-neutral premiums for private insurance are mandatory since December 21st 2012 for all insurance policies. In contrast, the Japanese automobile insurance market was deregulated in 1998 such that bisex tariffs were, to a certain extend, reintroduced to the market (Saito, 2006). Unisex tariffs are thus a continuing issue of policy debates in insurance markets globally. Nevertheless, even though such regulation is discussed often, there is very limited empirical evidence on its economic consequences.

In this paper, we provide evidence on whether rate regulation in insurance pricing leads to a change in the demand for insurance, specifically in the demand for guarantees in variable annuities. Even though other studies touched upon this issue before (Saito, 2006), we are the first who are able to take advantage of a natural experiment to analyze it. We are also the first to provide evidence on the effect of rate regulation in the annuity market. Other studies in this market suggest individuals to behave according to economic theory most of the time, but not always. Milevsky and Kyrychenko (2008) show that individuals with a guaranteed minimum benefit in their variable annuity contract choose a riskier portfolio than individuals without such a guarantee and are therefore behaving according to theoretical predictions. Similarly, Einav et al. (2010) show that people whose private information suggests them to have a higher mortality rate choose a higher annuity guarantee period than others, since their premium is

[1] An interesting point to this regard is raised by Finkelstein et al. (2009). They show that natural market reactions to the introduction of unisex tariffs will lead to contract designs which will prohibit effective redistribution between genders to a certain degree.

comparatively lower. In contrast, Knoller et al. (2014) show that while policyholders do react to the value of financial options and guarantees provided in their variable annuity contracts, their behavior is not always optimal.

Rate regulation has most commonly been criticized on the basis of adverse selection. Empirical evidence for adverse selection independent of rate regulation has been documented for the annuity market (e.g. Mitchell and McCarthy, 2002; Finkelstein and Poterba, 2004). The resulting welfare losses are substantial, as shown by Palmon and Spivak (2007) or Einav et al. (2010). An argument which connects rate regulation and adverse selection must, however, be based on the assumption that a change in premium leads to a change in insurance demand according to a negative price elasticity. Empirical evidence does not necessarily support this claim. Saito (2006) considers the heavily regulated Japanese automobile insurance market and finds no difference in coverage levels between different risk classes. He thus concludes that no causal effect of rate regulation and adverse selection or moral hazard can be found in his data. However, since he considers a static regulatory environment, no directly causal inferences can be made.

In our analysis, we use data from a large European life insurance company's portfolio of variable annuities from 2011 to 2014. Due to the unique regulatory conditions in the German annuity market, the German portion of our dataset contains a natural experiment to investigate the effect of unisex tariffs. While a share of the insurance policies sold prior to 2013 was priced on a bisex basis, so-called Riester-contracts were priced gender neutrally for the entire period of observation. As such, we can observe the change in individual choices when the pricing formula changes while controlling for other effects with the help of the Riester-contracts as a control group. Our analysis focuses on the initial choice of underlying portfolio and annuity guarantee period for the variable annuity. It thus also adds to the general literature on determinants of portfolio choice in retirement plans (e.g. Sundén and Surette, 1998; Agnew et al., 2003) and on portfolio choice in general (e.g. Frijns et al., 2008).

Independent of the general implications of our analysis for rate regulation in insurance markets, we are the first to analyze the effects of the E.U. Gender Directive empirically. Prior studies were limited to theoretical analyses (von Gaudecker and Weber, 2006) or the use of pre-regulation data (Aseervatham et al., 2013). However, with this analysis, we hope to provide a guideline to policymakers on the quantitative impact of unisex regulations globally.

Similar to other studies, we cannot directly observe the choice for or against annuities in our sample, but only those individuals which actually purchase a contract. However, as others have done before us, we can observe the choices which individuals have made within their contracts (Einav et al., 2010). Our study thus also ties into the literature of observing insurance demand through contract choices. This technique has been applied in several recent studies. Cohen and Einav (2007), Sydnor (2010) and Barseghyan et al. (2013) use it to estimate preference functionals of individuals. Other authors are more interested in behavioral effects like the demand for insurance against low probability high impact risks (Browne et al., 2013) or inertia in insurance choices (Handel, 2013).

The variable annuity market is particularly interesting in this regard, as it allows for more endogenous contract choices than traditional annuity products. In the specific contracts which we analyze, individuals buy a unit-linked variable annuity product that includes a guaranteed minimum income benefit (GMIB) as downside risk protection and the option for an annuity guarantee period (AGP). The insureds are able to choose the riskiness of the underlying portfolio. With more risk in the portfolio, the coverage due to the GMIB increases. Individuals who have a longer life expectancy have a higher expected benefit from this coverage and thus have to pay a higher premium for it. Since women have a higher life expectancy than men, they are ceteris paribus required to pay a higher premium for the GMIB than men in bisex tariffs. With the implementation of unisex tariffs, their GMIB premium is thus expected to fall while that for men is expected to rise.

Additionally, insureds choose the length of the AGP. This choice determines the minimum length of the period in which the annuity is paid out. If the insured dies before this period is expired, the rest of the payments guaranteed by the AGP are paid into his estate. As such, individuals with a longer life expectancy have a lower expected benefit from this coverage and thus have to pay a lower premium. The implementation of unisex tariffs will thus lead to opposite effects for the GMIB premium and the AGP premium. We thus expect opposite reactions to the change from bisex contracts to unisex contracts in the two choices for both genders. We formalize this hypothesis with an abstracted model below.

We find that both men and women choose a significantly lower risk level as their investment strategy in non-Riester contracts when the regulatory regime shifts from bisex to unisex contracts. Similarly, both genders choose a lower average annuity guarantee period due to the regulatory change. These effects persevere when using the Riester-contracts as a control group.

This implies a causal effect of unisex tariffs on both choices. Our results thus show that the introduction of unisex tariffs reduces the demand of individuals to take advantage of two important feature of variable annuities: participation in rising stock markets without the risk of losing the investment and guaranteed livelong annual consumption with an upheld bequest in case of an early death. Therefore, the change in policyholder's behavior induced by unisex tariffs leads to decreased consumer welfare from variable annuities.

The observed effect is equal for men and for women in both choices, which makes us unable to explain our results with classic economic theory. We thus do not find empirical support for the hypotheses derived from our model. We provide an alternative explanation of our findings based on the public perception of unisex tariffs. However, we do not have sufficient data to test this hypothesis and thus leave it open for future research.

After the introduction, the paper structure is as follows: In the next section, we describe our data and the natural experiment setting in detail. In the third section, we present a simple model that provides hypotheses about the change in contract choices due to rate regulation. We present our empirical strategy to test these hypotheses in the fourth section and show the results of our estimation. This section also contains robustness checks of our results. In section five, we discuss a potential explanation for our finding that men and women react equally to unisex contracts. The paper ends with some concluding remarks.

2 The Data

2.1 Contract Choices

Variable Annuities are unit-linked annuity contracts with one or more guaranteed minimum benefits.[2] For our entire analysis, we will focus on deferred annuities. This means that prior to the collection period of the annuity, there is a period of regular premium payments which can often be quite lengthy. The duration of the contract, i.e. the time between commencement date and maturity of the contract is often longer than 30 years. The premium payments are continuously invested into funds, the returns of which are accumulated throughout the saving period. Unlike in traditional unit-linked insurance plans, where the policyholder bears the entire financial risk of the returns, the minimum guarantees in variable annuities provide a downside risk protection.

[2]For a detailed overview of variable annuities see Bauer et al. (2008).

Our analysis focuses on products with a Guaranteed Minimum Income Benefit (GMIB) and the possibility for an annuity guarantee period (AGP). These products work as follows: The savings component of the premium is invested periodically, e.g. on a monthly basis, into managed funds. At commencement date, policyholders choose between three funds and therefore determine the risk-return profile of the investment. They can choose between a low risk fund with 30% stocks and 70% bonds as a target value, a medium risk fund with equal shares and a high risk one with 70% stocks and 30% bonds.

At maturity, the periodic annuity payment resulting from the annuitized fund value is compared to the GMIB, a guaranteed minimum annuity payment, and the higher value is paid out from then onwards. Once the annuity payment is determined, it is fixed and therefore no longer under financial risk. The financial risk resulting from the choice of the fund strategy is thus completely realized at maturity of the contract. The GMIB is known to the customer at commencement date and depends on the savings premium, duration until maturity and the customer's life expectancy. It does not change over the duration of the contract, unless the customer carries out contractual amendments. For this downside protection, a guarantee fee has to be paid that differs with the fund choice. This is because the share of stocks influences the volatility of the portfolio, which changes the extend to which the guarantee has to take effect. If gender is included in the pricing formula, the guarantee fee also differs between men and women. In this case, the fee is cheaper for men, because men have a shorter life expectancy. This means in case the GMIB has to be paid from maturity on, it does not have to be paid as long as for women.

In models of insurance markets, economic agents are usually assumed to have a fixed risk and choose their insurance coverage. In our GMIB setting, however, agents have a fixed level of coverage (the guarantee level) and choose their portfolio composition. In a certain sense, they thus choose their risk instead of their level of insurance coverage.[3] Nevertheless, the choice still influences the insurance coverage as is illustrated in Figure 1. The two panels show the possible range of the fund value (FV) across time in a stylized fashion. The development of a low risk portfolio (panel (a)) has a smaller spread than that of the high risk portfolio (panel (b)). Since the guaranteed interest rate through the GMIB is constant for all levels of risk, it is obvious, that the expected loss amount covered by it (the shaded area) increases with the riskiness of the

[3]To be precise, it is the choice of the size of the loss that is endogenous. The distribution of the loss, i.e. the performance of the risky asset, remains exogenous. The individual can only choose the share of stocks in his portfolio and therefore determine the outcomes.

portfolio for any contract duration t. As such, the choice of portfolio risk is also one of insurance coverage in the form of downside risk protection.

[Figure 1 about here]

Aside from the risk of the underlying portfolio, insureds can also choose their annuity guarantee period at commencement date. It can vary between 0 and 30 years.[4] Without an AGP, the periodic annuity payment ends with the death of the insured person. An AGP provides a guaranteed period over which the annuity will be paid, even if the insured person dies within this period. An annuity guarantee period therefore only makes sense if the policyholder has a bequest motive. Is the insured person still alive at the end of this guarantee period, the annuity payments stop with the death of the insured person.

As shown in Figure 2, the amount of coverage provided by a fixed AGP differs in the life expectancy of the insured. Panel (a) shows a stylized probability density function of a male with a given AGP while panel (b) shows a similar picture for women with the same AGP but a higher average life expectancy. As can be seen by comparing the two panels, the probability of being covered by the AGP (shaded area) is higher when the life expectancy is lower. As such, the fee for the AGP will be lower for women then for men when bisex tariffs are calculated.

[Figure 2 about here]

Other than the choice of the fund strategy and AGP, policyholders also have the typical choices in contract design that are known from traditional annuity products. They can choose the maturity of the contract, how much premium they want to contribute every year and whether to pay the premium in monthly or annual installments. We are able to observe all of these choice variables.

2.2 Natural Experiment Setting

We use data from a large European life insurer. Since we are taking advantage of the unique regulatory situation in Germany, we only use the German portfolio for our analysis. The dataset covers the period 2011 to 2014 and contains 18,764 observations.[5] The unit of observation is the contract and includes information about the choice of the underlying fund, the annuity

[4] Technically, the AGP can last until age 90 of the insured. In our data this commonly implies a maximum AGP of 30 years or less.

[5] Due to confidentiality issues we use a high quantile random sample of the original dataset to conceal the actual portfolio size.

guarantee period, several contract characteristics such as the premium, and the demographic characteristics age and gender (whether the contract was bought pre- or post regulation). Table 1 provides an overview of all variables currently used in the analysis.

[Table 1 about here]

The product is sold in five different versions with only small differences in the pricing. Besides the "regular" version – private insurance which is open to everybody – it is also distributed in three different ways of voluntary occupational pension insurance. In this case, the employer provides an annuity payment to his employees starting at retirement. These contracts are thus only open to people which are in employment at the commencement date. There are several ways how these payments can be provided with the help of an insurer. Our data includes Direct Insurance, which is the most frequent way, as well as Support Fund and Direct Grant. The difference to the regular version of the product, besides some minor pricing differences, is that employee contributions are tax deductible.

The last possible product category are so-called Riester contracts.[6] These contracts have a slightly higher administration fee than the other products, but the premium payments are partially subsidized by the German government.[7] The exact level of the subsidy is dependent on the number of children of the policyholder. The Riester contracts have been regulated by the German federal government to be priced unisex since 2006 and therefore over our entire observation period.

The unique combination of almost completely similar annuity products underlying different pricing regulation comprises a natural experiment with regard to the demand effects of unisex pricing. All non-Riester products were affected by the E.U. Gender Directive and were thus switched from bisex pricing to unisex pricing on December 21st 2012. This does not mean that contracts sold before December 21st 2012 were changed in the pricing, but rather that all contracts sold from this date on had to be priced unisex. Since the Riester product was priced in a unisex regime for the entire observation period, it was unaffected by the European legislation, i.e. there is no pricing difference between Riester contracts sold before and after the regulatory change. As such, they can serve as a control group for our analysis of the non-Riester contracts.

[6] For a detailed overview of the design of Riester contracts see Börsch-Supan and Wilke (2004) or von Gaudecker and Weber (2006).

[7] Furthermore, there are minor differences in the calculation of the GMIB between Riester and non-Riester contracts. However, these differences are of no consequence, since the Riester contracts only serve as a control group for potential changes in risk attitude or capital market expectations of new policyholders over time.

We provide an overview of the sample size pre- and post-regulation for both the treatment and the control group in Table 2.

[Table 2 about here]

Similar to several other studies before us, we consider the choices made by policyholders within a contract instead of the decision whether or not to purchase a contract at all (e.g. Cohen and Einav, 2007). Considering our set-up, it would also be possible to compare the shares of women in the treatment group and the control group before and after the regulatory intervention. However, we have reservations regarding such an analysis. Since we do not observe any customers who do not purchase the variable annuities, we cannot with certainty make any deductions regarding the behavior of the population. Riester contracts are generally seen as sensible if the insured is of low income and in full-time employment. The change in gender roles in Germany over time as well as the substantial development in the income of women and the increase of self-employed women over time would all drive an increase of the share of women in the treatment group even if no demand reaction was present.[8] Such a trend would bias our results and without being able to observe the non-purchasing policyholders, we would have no way to control for it. Furthermore, annuities sales are very reliant on the intermediary. If the market environment or the incentives for the intermediary change, a trend in the population can easily be observed. However, in the annuity product of our sample, there are no incentives for the intermediary to sell a specific annuity guarantee period or a specific risk level. The compensation of the intermediary is independent of the contract choices of the customer. As such, when only observing contract choices, our observations are free from any influence of the intermediary structure which might bias our results.

Tables 3 and 4 summarize the average share of risky assets and the average AGP length in the portfolio choices of new contracts pre- and post-regulation for both the treatment and the control group. Preliminary analysis points towards a significant effect of unisex tariffs on the demand for guarantees in variable annuities. When comparing the share of risky assets chosen in the different groups pre- and post-regulation, it is evident that while the share of risky assets in the Riester contracts increased, it decreased in all other contracts. This points towards a negative effect of unisex tariffs on the share of risky assets in the portfolio underlying variable annuities.

[8] As an example for this trend, the income of women in Germany grew by 9.41% between 2008 and 2012, while that of men grew by only 6.91%.

[Table 3 about here]

A similar observation can be made with regard to the annuity guarantee period. While there is a slight increase in the average AGP chosen before and after the regulatory change in the non-Riester contracts, the increase in the Riester contracts is about six times as large. This suggests that even though the time trend is positive from before 2012 to after 2012, the regulation implies a downward shift in demand for an annuity guarantee period.

[Table 4 about here]

In the following section, we will develop a more thorough theoretical model than the qualitative argument presented in section 2.1 above. We will use it to derive predictions for the effect of rate regulation on the choice of portfolio risk and AGP. We will then test the derived hypotheses in a more stringent empirical setting than descriptive statistics in section 4.

3 Model

To provide a rigorous link between price changes, the choice of portfolio risk, the choice of the annuity guarantee period and the different risk types (males and females) in the specific setting of our data, we use a very simple model of a variable annuity contract. In the model, an agent of type $i \in \{m, f\}$ lives to a maximum of two periods. The respective utility functions for each of the periods are denoted $U_1(\cdot)$ and $U_2(\cdot)$ and display risk aversion.[9]

In the first period, the agent is alive for certain and has wealth w. The agent is still alive in period t_2 with a survival probability κ_i, and has no source of income in that period. If the individual does not survive to t_2, his utility of bequest is measured by the non-decreasing, concave function $U_B(\cdot)$. The idea behind this bequest function is that the individual not only benefits from lifetime consumption, but also from inheritance to his descendants. This concept goes back to Yaari (1965) and was further analyzed by Fischer (1973), Hakansson (1969) and Campbell (1980). Following Abel (1986), we assume that men and women differ in their respective survival probability κ_i.[10] $\kappa_f > \kappa_m$ reflects the fact that women have a higher life expectancy than men.

[9]This approach includes the special case of the discounted expected utility model, where utility in the second period is given by $\delta U_1(\cdot)$, with $\delta > 0$ being the discount factor associated with the pure rate of time preference. Note, however, that the approach utilized here is more comprehensive than that. It allows for differing risk aversion and/or shape of the utility function between the two periods.

[10]Since Abel (1986), the concept of heterogenous survival probabilities has been used in several studies to analyze adverse selection in the annuity market, see Brugiavini (1993), Brown (2001) and Steinorth (2012), just to name a few.

The model abstracts in the sense that there is usually not much difference between men and women in the probability of reaching maturity of an annuity contract, but rather a difference in the expected length of the annuity payment period afterwards. Nevertheless, the intuition is still carried over in our model. One could also add periods in between our two periods. Since these would be the same for both agents, they would not influence our results. In the interest of brevity of both the model and this study, we will use the abstract setting of two periods.

The only mode to transfer wealth between t_1 and t_2 is the variable annuity. The amount invested in this annuity is denoted I and is fixed in advance. The agent has made an exogenous choice of investing in the VA and now has to make the choice of portfolio composition. The annuity amount is invested into a risky and a risk-free asset, which render returns $\tilde{z}$ and r, respectively. We assume $\int z \, \mathrm{d}F(z) > r$. This is not a particularly restrictive assumption since for any risk averse decision-maker, $\int z \, \mathrm{d}F(z) \leq r$ would imply the sole purchase of the risk free asset. The percentage of the portfolio invested in the risky asset is denoted c_i and we do not place any restrictions on it (i.e. short selling and lending are allowed without transaction costs). However, due to $\int z \, \mathrm{d}F(z) > r$ it is obvious, that c_i will never be negative since such a solution would always be first order stochastically dominated by a positive c_i of the same absolute value.

The VA includes a GMIB clause which means it renders at least return g independent of the portfolio composition and the realization of $\tilde{z}$. Therefore, the return of the VA in case of survival until period t_2 can be written as $max\{c_i\tilde{z} + (1 - c_i)r; g\}$. For simplification we assume $g = r$ which renders $Pr(c_i\tilde{z} + (1 - c_i)r < g) = Pr(\tilde{z} < r)$.[11] We abbreviate $Pr(\tilde{z} \geq r)$ to p, and $1 + c_i\tilde{z} + (1 - c_i)r$ to $\widetilde{\gamma(c)}$. The GMIB is prized actuarially fair with premium

$$\pi_c(c_i, \kappa_i) = \kappa_i c_i (1 - p) \left(r - \int_{\tilde{z}<r} z \, \mathrm{d}F(z) \right) I.$$

The GMIB premium is paid to insure the downside risk, i.e. the difference between the realization of $\tilde{z}$ and the guaranteed return r. This equation and the fact that the pay-off in case of survival is equal to $max\{c_i\tilde{z} + (1 - c_i)r; g\}I$ shows the first of two major simplifications we make in the model. From the perspective of the insurance company, the downside risk protection is prized actuarially fair. However, the annuity is not. Instead, there is no compensation in case of death coming from the initial investment I. This is owed to the fact that our data does not let us observe the non purchase of annuities or which other transactions the insured makes to

[11] This simplification does not have a large influence when considering that g in or dataset is oriented towards the r that can be achieved in the market.

save for retirement. For an actuarially fair annuity, the pay-out would have to be modeled as $I\kappa_i^{-1}$ or the investment would have to be reduced to $\kappa_i I$. However, changing the regulatory framework and thus the κ_i used in the pricing of the annuity, would change the benefit of investing in the annuity without allowing the decision-maker to change the initial investment. If we would endogenize I, we would make predictions which are difficult to answer without observing the other transactions of the individual. Since we want to focus on the choices made within the contract and not on the contract itself, we use this abstracted setting in the interest of consistency between model and empirical set-up.

Additionally, the agent has the option to include an annuity guarantee period into his VA contract. The AGP is modeled as simple as possible. The agent chooses some amount a_i in advance. If he does not survive to the second period, this amount will be paid into his estate. The premium is again paid in the first period and expected value neutral. The AGP is prized actuarially fair with premium

$$\pi_a(a_i, \kappa_i) = (1 - \kappa_i)a_i$$

This shows the second major simplification used in this model. Instead of a time period, we model the annuity guarantee period as a lump sum payment a_i. a_i represents the net present value of future guaranteed annuity payments and is paid into the individual's estate in case of death in t_2. The reason for this abstraction is of technical nature. In a theoretical model with annuity guarantee as a period of time individuals could substitute a_i with c_i and yield the same expected utility. A shorter annuity guarantee period could be compensated with a higher share of stocks and therefore higher (expected) returns per period in case of untimely death. However, in reality, policyholders typically want to protect their dependents by providing them an ongoing stream of income in case of premature death. Thereby, especially the time of protection and not so much the possible higher returns matter for this. Modeling a lump sum annuity guarantee reflects this choice situation more realistic.

The agent is assumed to maximize his expected utility of lifetime consumption and bequest denoted as V^{κ_i}, the objective function reads:

$$\begin{aligned} V^{\kappa_i} =& U_1(w_1 - I - \pi_c(c_i, \kappa_i) - \pi_a(a_i, \kappa_i)) + \kappa_i(1-p)U_2((1+r)I) \\ &+ \kappa_i p \int_{\tilde{z}>r} U_2(\gamma(c_i)I)\,\mathrm{d}F(z) + (1-\kappa_i)U_B(a_i) \end{aligned}$$

Partial derivatives are denoted by subscripts of model parameters. For example, $V_{c_i}^{\kappa_i}$ stands for $\partial V^{\kappa_i}/\partial c_i$. We use this notation throughout the paper. The optimal risk and AGP choices $(c_i^{\kappa_i}, a_i^{\kappa_i})$ are characterized by the first order conditions [12]

$$\begin{aligned}
V_c^{\kappa_i} =& -\kappa_i(1-p)\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right) IU_1'(w-I-\pi_c(c_i,\kappa_i)-\pi_a(a_i,\kappa_i)) \\
&+\kappa_i p\int_{\tilde{z}>r}\frac{\partial\gamma}{\partial c_i} IU_2(\gamma(c_i)I)\,\mathrm{d}F(z)=0 \\
V_a^{\kappa_i} =& -(1-\kappa_i)U_1'(w-I-\pi_c(c_i,\kappa_i)-\pi_a(a_i,\kappa_i))+(1-\kappa_i)U_B'(a_i)=0.
\end{aligned}$$

These conditions show the common trade-off of marginal benefit and marginal cost, i.e. the costs occurred by the insurance premiums $\pi_c(c,\kappa_i)$ and $\pi_a(a,\kappa_i)$ in the first period are contrasted with the benefits of the respective insurance in the second period. It is noteworthy that the marginal utility of increasing c_i is zero in case of a detrimental development of the risky asset. Since the GMIB protects the agent from a bad outcome of the random variable $\tilde{z}$, the marginal effect in the second period is only positive.

In the unisex regime, the pricing of the insurance contracts changes. Instead of using individual survival probabilities for each agent type, the insurance company is now forced to use a common survival probability $\overline{\kappa}$ instead. Several possible scenarios exist on how the pricing could be organized. A regulatory system in which the risk type cannot be used for pricing insurance policies is hypothesized to lead to adverse selection. Economic intuition would thus suggest the insurer to price their policies such that either a separating equilibrium (Rothschild and Stiglitz, 1976) or a pooling equilibrium (e.g. Wilson, 1977) obtains. The former would imply a non linear

[12] Note that, due to the concavity of V^{κ_i}, $(c^{\kappa_i}, a^{\kappa_i})$ maximizes expected utility, as

$$\begin{aligned}
V_{cc}^{\kappa_i} =& \kappa_i^2(1-p)^2\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right)^2 I^2U_1''(w-I-\pi_c(c_i,\kappa_i)-\pi_a(a_i,\kappa_i)) \\
&+\kappa_i p\int_{\tilde{z}>r}(\frac{\partial\gamma}{\partial c_i})^2 I^2U_2(\gamma(c_i)I)\,\mathrm{d}F(z)<0 \\
V_{aa}^{\kappa_i} =& (1-\kappa_i)^2U_1''(w-I-\pi_c(c_i,\kappa_i)-\pi_a(a_i,\kappa_i))+(1-\kappa_i)U_B''(a_i)<0
\end{aligned}$$

and the determinant of the Hessian is positive, i.e.

$$\begin{aligned}
V_{cc}^{\kappa_i}V_{aa}^{\kappa_i}-(V_{ca}^{\kappa_i})^2 =& \kappa_i^2(1-p)^2\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right)^2 I^2(1-\kappa_i)U_1''(w-I-\pi_c(c_i,\kappa_i)-\pi_a(a_i,\kappa_i))U_B''(a_i) \\
&+\kappa_i p(1-\kappa_i)^2U_1''(w-I-\pi_c(c_i,\kappa_i)-\pi_a(a_i,\kappa_i))\int_{\tilde{z}>r}(\frac{\partial\gamma}{\partial c_i})^2 I^2U_2(\gamma(c_i)I)\,\mathrm{d}F(z) \\
&+\kappa_i p(1-\kappa_i)\int_{\tilde{z}>r}(\frac{\partial\gamma}{\partial c_i})^2 I^2U_2(\gamma(c_i)I)\,\mathrm{d}F(z)U_B''(a_i) \\
&>0.
\end{aligned}$$

change in the survival probability used for pricing. For high levels of coverage (high c_i and high a_i), the respective $\overline{\kappa}$'s would be chosen such that they are close to that of the high risk groups (females and males, respectively). This would ensure self selection. In a pooling equilibrium, the insurance company would be interested in offering only one contract. In our setting, this would be made feasible by prizing one contract actuarially fair and overpricing all other contracts such that they are unattractive to the customers.

The insurance company which provided our data uses neither of the two approaches. Instead, they use a linear interpolation between the two survival probabilities. This corresponds to the equilibrium concept developed by Arrow (1970), which was analyzed in detail by Pauly (1974) and Schmalensee (1984). In this so called linear-pricing equilibrium the different risk types pay the same constant price per unit of coverage, which yields pooling contracts. Therefore there is no price discrimination. Both genders pay the same price per unit of coverage for the two contract choices, and any additional amount of coverage has the same price as the first amount of coverage. Furthermore, this means that there is no quantity rationing, i.e. the levels of purchasable coverage do not change compared to the full information setting. In the interest of confidentiality, we will not go into detail regarding the new pricing approach. For our purpose it is sufficient to say that $\kappa_f > \overline{\kappa} > \kappa_m$ and that $\overline{\kappa}$ is fixed for all possible contract choices. The objective function now reads

$$\begin{aligned} V^{\overline{\kappa}} =& U_1(w - I - \pi_c(c_i, \overline{\kappa}) - \pi_a(a_i, \overline{\kappa})) + \kappa_i(1-p)U_2((1+g)I) \\ &+ \kappa_i p \int_{\tilde{z}>r} U_2(\gamma(c_i)I)\,\mathrm{d}F(z) + (1-\kappa_i)U_B(a_i) \end{aligned}$$

Individuals now choose the contract parameters $(c_i^{\overline{\kappa}}, a_i^{\overline{\kappa}})$ that maximize expected utility at the pooling price. The individual contract parameter choice is observable by the insurer and is used to determine the premium. The price only depends on the individual level of coverage, not on the individual survival probability. Thus, low risks subsidize high risks.[13] As we are

[13]Note that it is not unambiguously clear if men or women are the high risk. For the downside protection with the pooling premium

$$\pi_c(c_i, \overline{\kappa}) = \overline{\kappa} c_i (1-p) \left(r - \int_{\tilde{z}<r} z\,\mathrm{d}F(z) \right) I.$$

women are the high risks and get subsidized by men, as $\kappa_f > \overline{\kappa} > \kappa_m$. However, for the annuity guarantee with pooling premium

$$\pi_a(a_i, \overline{\kappa}) = (1-\overline{\kappa}) a_i$$

men are the high risks and get subsidized by women. Men and women therefore cross-subsidize each other. Fluet and Pannequin (1997) provide a detailed discussion and a general model of asymmetric information in multiple antithetic risk contracts. Distinct from our model, they assume a separating equilibrium.

interested in the change in individual choices, we look at $\partial c_i/\partial\overline{\kappa}$ and $\partial a_i/\partial\overline{\kappa}$, i.e. we analyze how the optimal choices of risk and AGP $(c_i^{\overline{\kappa}}, a_i^{\overline{\kappa}})$ vary if $\overline{\kappa}$ changes.

The following proposition shows the effect of unisex pricing on the optimal choice of downside protection c_i and AGP a_i. It differentiates between the cases $I > I^T$ and $I < I^T$, with

$$I^T := \frac{a_i}{c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)}$$

displaying the ratio of annuity guarantee risk and portfolio downside risk. Using comparative statics, we can obtain:

Proposition 1. *The change in product choices due to the unisex regime can be separated into two cases depending on I:*

i) If I is larger than I^T, the share of the risky portfolio is decreasing in $\overline{\kappa}$ and it is a sufficient condition for the annuity guarantee period to be increasing in $\overline{\kappa}$ that the absolute risk aversion in the first period is not too high.

ii) If I is smaller than I^T, the annuity guarantee period is increasing in $\overline{\kappa}$ and it is a sufficient condition for the share of the risky portfolio to be decreasing in $\overline{\kappa}$ that the absolute risk aversion in the first period is not too high.

Proof. See Appendix □

Irrespective of the value of I, we always have at least one of the choice variables reacting to rate regulation as was argued in section 2.1. However, there is also always one choice variable for which the picture is not so clear. This is due to the fact that there are two competing effects. On the one hand, there is the direct effect as described in section 2.1. A change in $\overline{\kappa}$ affects the premiums π_c and π_a that have to be paid in the first period and therefore the demand for c_i and a_i, respectively. But on the other hand, there is also a negative substitution effect, as the change in $\overline{\kappa}$ affects both premiums at the same time. Therefore a change in the premium of one risk might confound period-1 consumption so much that the individual has to compensate this with a reverse choice of the other risk. Conditions (6) and (5) determine when the substitution effect prevails.[14]

Although these conditions are very complex, they are not very restrictive. The thresholds are an upper limit for the individual's absolute risk aversion. They depend on several parameters,

[14]To make this point clearer, we provide an example and a detailed analysis of the conditions in the Appendix.

but might indeed often lie above 1. However, common values for the coefficient of absolute risk aversion are usually measured to be well below 0.1 (Feldman and Dowd, 1991). Combining this with the fact that the conditions are merely sufficient, but not necessary for the comparative statics to behave in this fashion, one can argue that our proposition applies to the vast majority of cases.

As Proposition 1 gives one universally valid prediction for each choice variable and one prediction that holds in the majority of cases, we can derive empirically testable hypotheses from the model. We now only need to interpret Proposition 1 in light of the fact that in the unisex tariffs $\kappa_f > \overline{\kappa} > \kappa_m$.

Hypothesis 1. *In the unisex tariffs, men will choose a lower share of risky assets in their portfolio than in the bisex tariffs.*

Hypothesis 2. *In the unisex tariffs, women will choose a higher share of risky assets in their portfolio than in the bisex tariffs.*

Hypothesis 3. *In the unisex tariffs, men will choose a longer annuity guarantee period than in the bisex tariffs.*

Hypothesis 4. *In the unisex tariffs, women will choose a shorter annuity guarantee period than in the bisex tariffs.*

The next section develops an empirical strategy to test these predictions and reports the results.

4 Estimation and Results

4.1 Empirical Methodology

To examine the choice of portfolio composition and the annuity guarantee period in our data, we take advantage of the natural experiment setting and use a difference in difference estimation. The first endogenous choice variable in our data, the riskiness of the portfolio, is a categorical variable in nature. This would point towards using an ordered probit estimation for the evaluation of the effects. However, since we make heavy use of interaction effects, we utilize ordinary least squares regression instead. Interaction effects are hard to interpret in non-linear models (Ai and Norton, 2003). We thus make the simplifying assumption that the difference in

choice between the low risk portfolio and the medium risk portfolio is equal to that between the medium risk portfolio and the high risk portfolio.[15] The second endogenous choice variable, the AGP, is a discrete variable with a large number of categories and will thus be treated as if it was continuous.

We denote our coefficients of interest by $\beta_k^{(c,a)}$ and use $\vec{C}$ as a vector of control variables and $\vec{\gamma}^{(c,a)}$ as their coefficients. The control variables also include dummies which are coded to imply the current contract generation at the time the contract was signed. Different generations have slight differences in pricing, but no differences in terms of the guarantees which can be chosen within a contract. The omitted category is the last contract generation which comprises over a third of the total contracts observed and was put into force about a year before the unisex implementation. We provide a robustness check regarding this dummy specification in the next section.

The superindex of a coefficient indicates whether the coefficient is used in the estimation of portfolio risk, or in the estimation of the AGP. Our system of difference in difference estimations looks as follows[16]:

$$\begin{aligned} risk = &\beta_0^c + \beta_1^c t_{\geq 21.12.2012} + \beta_2^c treat + \beta_3^c t_{\geq 21.12.2012} \times treat + \beta_4^c female + \beta_5^c t_{\geq 21.12.2012} \times female \\ &+ \beta_6^c treat \times female + \beta_7^c t_{\geq 21.12.2012} \times treat \times female + \vec{\gamma}^c \vec{C} + \epsilon^c \end{aligned} \tag{1}$$

$$\begin{aligned} agp = &\beta_0^a + \beta_1^a t_{\geq 21.12.2012} + \beta_2^a treat + \beta_3^a t_{\geq 21.12.2012} \times treat + \beta_4^a female + \beta_5^a t_{\geq 21.12.2012} \times female \\ &+ \beta_6^a treat \times female + \beta_7^a t_{\geq 21.12.2012} \times treat \times female + \vec{\gamma}^a \vec{C} + \epsilon^a \end{aligned} \tag{2}$$

Our theoretical model implies that there is a relationship between optimal choices of a_i and c_i. Furthermore, it is imaginable that bequest motive and risk aversion have a statistical relationship. In such a case, there would also be a correlation between the two choices. To allow for such an interdependence, we could use a seemingly unrelated regression model in our estimation. This would allow the error terms ϵ^c and ϵ^a to be correlated by some coefficient ρ. However, due to the fact that our vectors of regressors in both equations are equal, there is no

[15] While this assumption is unproblematic when looking at the shares of stocks in the different portfolios, it can be questioned from the perspective of the utility differences between the different portfolios. However, for the ease of interpretation, we remain with using the linear model. Results from a non-linear estimation show signs of coefficients and significance levels equal to the ones in our estimation as is reported in section 4.3. However, since the marginal effects of interaction effects in categorical estimation models are neither constant in magnitude nor in sign, this has to be seen as preliminary evidence.

[16] Standard errors are robust to allow for heteroscedasticity.

informational advantage from such a specification (Davidson and MacKinnon, 1993). We thus use two OLS estimations instead.

As listed in Table 1, the variable $t_{\geq 21.12.2012}$ indicates whether the unisex regulation was in effect at the commencement date of the annuity contract. Note that for this variable it is irrelevant whether the contract was actually affected by the regulation or not. The variable *treat* is coded to take the value one if the contract was in the treatment group. As such, the coefficient of the interaction effect of the two variables, β_3,[17] measures the difference in reaction to the unisex regulation between the two different groups of contracts. With this specification of only two periods of time, before or after December 21$^{\text{st}}$ 2012, we do not run into the problem of serial correlation in difference in difference estimations (Bertrand et al., 2004).

Our estimation differs from a common difference in difference estimation by the existence of the dummy coefficient for females as well as its interaction with all other relevant coefficients. In fact, when only estimating the model with the coefficients β_0 through β_3 (estimations (1) and (2) in Table 6), we have a regular difference in difference estimation. The dummy coefficient for female policyholders is introduced to differentiate the effects of unisex tariffs on men and women. This is necessary to test the hypotheses derived above. The coefficients β_5 through β_7 are used to identify these differences. Some guidance on the interpretation of all relevant coefficients in the estimation is given in Table 5. The table gives an overview which coefficients take effect for which group of the sample. For example, the risk taken by men in the control group prior to the regulation is only measured by β_0^c, while that of women in the treatment group after the regulation is measured by all eight β^c coefficients.

[Table 5 about here]

Based on these considerations, we can now link our hypotheses with the empirical model. When estimating the full model as indicated in equations (1) and (2), we would expect the coefficients β_3^c and β_3^a to have signs in accordance with the Hypotheses 1 and 3 since men are the omitted group. Since $\overline{\kappa} > \kappa_m$, we would thus expect $\beta_3^c < 0$ and $\beta_3^a > 0$. The extend to which women differ from men in both choices when the rate regulation is implemented is indicated by the two coefficients β_7^c and β_7^a. Since women are hypothesized to react in the opposite direction of men, we would not only expect $\beta_7^c > 0$ and $\beta_7^a < 0$ but rather $\beta_7^c > -\beta_3^c$ and $\beta_7^a < -\beta_3^a$.

[17] In the interest of legibility, we drop the superscript (c, a) when discussing the general case in the following.

4.2 Results

The results of our four main estimations are presented in Table 6. The table reports a total of four regressions. In the first two, simplified difference in difference estimations without interaction effects for women are reported. While the first estimation only includes those coefficients relevant for the hypotheses, the second estimation also includes the vector of control variables[18]. Estimations (3) and (4) include the entire specification given in equations (1) and (2). These estimations again differ with respect to the control vector.

[Table 6 about here]

The coefficient β_3 is significantly smaller than zero for both the annuity guarantee period and the risky asset choice. We thus find that unisex tariffs lead to generally lesser demand for guarantees within variable annuity contracts. This effect is robust to all four specifications, even though it is smaller when exogenous factors are controlled for. The marginal effect of the change from bisex tariffs to unisex tariffs is rather large. As we can see in our preferred specification, regression number four, the average share of risky assets declines by a more than three percentage points, which is about 8.5% of the entire possible range and about 7.9% of the average share of risky assets in the treatment group pre-regulation. The average annuity guarantee period declines by 0.829 years which is about 2.75% of the entire possible range and about 5.3% of the average AGP in the treatment group pre-regulation.

To see the significance of the effect of unisex tariffs, we can compare it to the gender effect in decisions under risk. Even though some debate exists (Schubert et al., 1999), it is generally accepted that women are more risk averse than men in at least some decision situations (Powell and Ansic, 1997; Jianakoplos and Bernasek, 1998; Halek and Eisenhauer, 2001). In correspondence with this result, we find that women in Riester contracts choose to allocate about 0.8% less to risky assets in their portfolio than men.[19] From this we can infer that the effect of unisex tariffs on risk taking in our data is about four times as large as the gender effect.

While our results show a statistically as well as economically significant effect of rate regulation in the form of unisex tariffs, the observed effects are not in line with some of our hypotheses. The coefficient β_3^a is negative instead of positive as was predicted by Hypothesis 3. This implies

[18] Throughout the paper, the vector of control variables always includes a linear and a squared term for *age* as the influence of age on insurance choices is often seen to be non-linear.

[19] If not stated otherwise, all our results refer to regression number four as it is the most comprehensive estimation.

that the demand by men for an annuity guarantee period declines even though selecting this contract feature is actually cheaper after the regulation than before. Similarly, we observe a non-significant and almost zero coefficient β_7^c. This implies that women also select a smaller share of risky assets after the unisex implementation. They thus behave similar to the men in the sample, even though the price for having a higher share of risky assets declines after the regulatory intervention. Our results thus also contradict Hypothesis 2.

Since β_3^a is negative, a non-significant coefficient β_7^a is actually in line with Hypothesis 4. Women buy less annuity guarantee period in the time after the unisex implementation than before. They thus react to the upward shift in the AGP premium as predicted. The reaction with regard to the choice of risky assets by men is also in line with our hypothesis, since β_3^c is negative as Hypothesis 1 predicted.

Even though only two of the hypotheses derived from our model were contradicted by the data, while the other two find statistical support, we nevertheless interpret the empirical results as a rejection of our model. We predicted that men and women would react in both choice situations as the price effects would predict. That means if the price of a contract feature would increase, people would buy less of it and if it would decrease people would buy more of it. However, our empirical results paint a different picture. We observe a universal downward shift in the demand for guarantees within variable annuities due to the rate regulation.

This result is puzzling from the perspective of traditional economics. It could thus raise the suspicion that our analysis might contain a bias which would explain such results. In the following section, we report robustness checks which alleviate this concern with regard to the econometric specification and selection bias.

4.3 Robustness

We start with examining the robustness of our results with respect to the econometric specification. A first question is whether the effect of the unisex regulation could be spurious because of different contract generations being examined in a single estimation. The dummy coefficients of the tariff generations should pick up any such spurious effects, but we nevertheless estimated our specification considering only the data from the newest contract generation. As would be expected when only a third of the data is used, our coefficients are less significant than when the full dataset is considered (the t-values for β_3^c and β_3^a are 1.956 and 1.646, respectively). Nevertheless, the coefficients are equal in sign. The marginal effect of the rate regulation on the

share of risky asset is lower, but still at slightly less than 1.9%. The marginal effect on the AGP choice has an absolute value of 0.678 and is thus almost equal to that for the full dataset. The coefficients β_7^c and β_7^a remain close to zero and insignificant. The complete results of regression four when estimated only with the newest generation of tariffs is given as estimation (5) in Table 7.

[Table 7 about here]

A second possible problem of the specification could be the coding of the share of risky assets as an interval variable. We thus repeat the estimation for equation (1) with an ordered probit estimation both for the full sample and for the newest contract generation only. Results are reported as estimations (6) and (7) in Table 7. It is apparent that the results do not change in terms of sign and significance. However, since we interpret interaction effects when regarding coefficients β_3^c and β_7^c, the reservations of Ai and Norton (2003) apply. As such, we do not make any inference about the marginal effects and report estimations (6) and (7) as preliminary evidence only.

Two possible sample selection biases could apply to our data. We will cover both of them in order. The first possible bias is a difference in the risk attitude of women which buy the non-Riester contracts after December 21$^{\text{st}}$ 2012 and women which bought them before that date. Such a difference between the people who bought the contracts in the bisex regime and those who bought the contracts in the unisex regime could explain the downward shift in the demand for risky assets by women if the women buying the contract before the regulation were less risk averse than those buying it afterwards. Similarly, our results regarding the AGP could be explained by a difference in the bequest motive of men which buy the non-Riester contracts after December 21$^{\text{st}}$ 2012 and men who bought them before that date. The difference in difference estimation should pick up any changes in the general population of our sample. However, specific changes in the population of policyholders which buy non-Riester contracts could lead to a sample selection bias.

A possible explanation for a difference in characteristics for women would be that women who were particularly risk averse chose not to buy the variable annuities observed here, because their structure as a unit linked product was not as appealing as a traditional savings product without any risk in the pay-off. However, with the introduction of unisex products, the GMIB

downside protection became cheaper for women and thus the variable annuity more attractive for risk averse women.

A similar argument cannot be made regarding the bequest motive. Under the bisex tariffs, men have to pay a comparatively expensive premium for having an AGP. As such, men with a strong bequest motive would tend not to buy this type of variable annuity contract but would rather be attracted to other savings devices. Once the unisex tariffs are implemented and the AGP becomes cheaper, such men are more likely to buy a variable annuity. If anything, a selection bias would thus lead to an overestimation of an increased demand for the AGP. This is clearly not supported by our data.

As we can see in Table 8, the share of women in the treatment group increases significantly after the implementation of unisex tariffs. This in itself will not bias our results, except if the men and women considered in the treatment group have different characteristics than those individuals before December 21st 2012.

[Table 8 about here]

There are certain reasons why we deem such a selection effect to be unrealistic. The first is that while unisex tariffs might not have been available for products in our treatment group, Riester contracts were available on a unisex basis before the European legislation took effect. Thus, any risk averse women which could be moved by unisex tariffs to buy a unit-linked contract, could have done so before with a Riester contract.[20] The second reason why we do not think that a major difference exists between those individuals that buy non-Riester annuities before December 21st 2012 and those that buy them afterwards is that they do not differ on any of the observable characteristics. In our sample, neither age, duration, payment modalities nor the size of the annuity (that is, the annual premium) differ between these two groups if gender is controlled for.

We nevertheless conduct an empirical test that could tease out a sample selection bias. When looking at the two subgroups that comprise our treatment group, the regular contracts and the occupational pension insurance, we see that the relative increase of women in the population is larger in the occupational pension contracts (40.87%) than in the regular contracts (35.39%). Thus, any selection bias should be more pronounced in the former contracts than in the latter.

[20] Even though the Riester contracts in our sample have a slightly higher administrative fee than the other products, this difference in pricing is nullified by the subsidies from the German federal government. Thus, there should be no selection effects of the contracts on this basis.

However, when conducting our estimation for both groups separately, we can observe no such effects. We can see in the estimations (8) and (9) in Table 7 that the estimated coefficients of the three-way interaction term $t_{\geq 21.12.2012} \times treat \times female$ in the risk regression are small, not statistically different from zero in both estimations and do not differ from one another at any common level of statistical significance ($\chi^2_{(1)} = 0.6, n.s.$). This result suggests that no selection bias in the sense of a difference in the preference parameters of the treatment group pre- and post-regulation exists.

A second possible issue of our estimation could be an issue of endogenous treatment selection. Even though this is not true for all individuals in the data, some had the option of choosing whether to buy a Riester or a non-Riester contract. It could be imagined that this choice was affected by the implementation of unisex tariffs. As such, the regressor *treat* could be endogenous to the decision problem.

As a robustness check for such an endogeneity problem, we report a three stage least squares estimation with an endogenized treatment effect in Table 9. We use the estimation strategy proposed in Wooldridge (2010). It takes advantage of the binary nature of the endogenous variable *treat*. As any other instrumental variable estimation, we need to instrument for the choice of contract. We do this by using the information about the distribution channel. The argument for this instrument is as follows. Different distribution channels have different incentives for selling Riester contracts. This could be due to differences in the corporate strategy of the different types of intermediary or solely due to monetary incentives of the individual salesmen. However, no perceivable difference in incentives exist regarding the choices of the individual within a given contract. As such, there seems to be no direct influence of the distribution channel on our dependent variables of interest, making it a good instrument to use in the analysis.

Before the first stage estimation, we use a probit estimation to refine our instrument. Abbreviating the vector of distribution channel dummies as $\vec{d}$, we can write this estimation as:

$$Pr\left(treat = 1|\vec{x}, \vec{d}\right) = \Phi\left(\delta_0 + \delta_1 t_{\geq 21.12.2012} + \delta_2 female + \vec{\delta}_3\vec{C} + \vec{\delta}_4\vec{d} + u\right) \tag{3}$$

From this estimation, we use the predicted probabilities for choosing a contract from the treatment group and use them in interaction terms as we used the dummy variable *treat* above. This renders a vector of instrument $\vec{p}$. We then use this vector as instruments for *treat* and

all its interaction terms in the first stage estimation of a three stage least squares estimation of equations (1) and (2).

[Table 9 about here]

As can be seen in the table, our results remain almost unchanged in sign when endogenizing the treatment choice via an instrumental variable estimation. The IV structure makes the treatment and its interactions more significant and lets them carry stronger marginal effects. However, implications remain unchanged. The coefficient β_7^a remains insignificant and is even reduced in size. The coefficient β_7^c becomes more negative and slightly significant. However, this would only imply that women reduce their portfolio risk even further than men due to the regulatory change and would thus still imply a rejection of our theoretical model. Since our results become stronger when endogenizing treatment choice, we are left with the conclusion that endogenous treatment choice is not the reason that our results are contrary to traditional economic theory.

5 Discussion

Our results in combination with the variety of robustness checks which are reported in this paper suggest an effect of unisex tariffs on the portfolio choice in GMIB annuities which cannot be explained by our model. While this might constitute a "puzzle" from the perspective of traditional economics, there might be a relatively simple explanation when considering the media coverage of unisex tariffs in the years preceding the implementation of the E.U. Gender Directive.[21] Early opinions voiced by the association of German insurance companies (*Gesamtverband Deutscher Versicherungswirtschaft*, GDV) paint a negative picture for policyholders. The GDV issued a press release on September 9$^{\text{th}}$ 2012 in which it states that various policies might become more expensive for either men or women (Gesamtverband Deutscher Versicherungswirtschaft, 2012). Particularly when reading the statement superficially, consumers might gain the impression that there is no benefit to any group involved.[22] This impression might have been intensified by further statements from insurance company officials. For example, Walter Boterman, CEO of the

[21] All following quote were translated from German by the authors. Emphases are kept as in the original.

[22] The only passage in the statement which could relate to a better standing of a gender in a specific policy is immediately put into perspective: "Some insurance policies will have the tendency to become cheaper for women and others will have this tendency for men. However, on average both genders will be burdened due to reactions of customers and uncertainty premiums."

nationwide operating life insurance company Alte Leipziger, was quoted by the Sunday newspaper Die Zeit. He summarized the issue as follows: "We never said that Unisex is going to be cheap for anybody." (Die Zeit, 2012).

Similar sentiments can be seen in other sources of information at the time. Germany's most circulated tabloid Bild featured the headline: "Unisex Tariffs: Why now EVERYBODY pays more" (Bild, 2013). The federation of German consumer organisations (*Verbraucherzentrale Bundesverband*), an association funded by and acting on behalf of the German government, issued a press release before the introduction of unisex tariffs expressing their concern about "abusive premium increases" due to unisex tariffs (Verbraucherzentrale Bundesverband, 2011). Axel Kleinlein, CEO of the German Association of the Insured (*Bund der Versicherten*), a non-profit consumer protection organisation focused on insurance, was quoted on the issue in several widely circulated news outlets. He stated before the implementation that: "There is always one gender in these type of tariffs, for which it gets more expensive. But that does not mean that it gets cheaper for the others" (Die Welt, 2012). After the implementation, he evaluated it as "blatantly and crassly at the expense of the customers" (Der Focus, 2013).

It might thus have been the public opinion that unisex tariffs lead to a generally worse outcome for all involved. This could mean that those policyholders who were buying annuities after the implementation of the E.U. Gender Directive were following the implicit advice of the media and interested parties to purchase less GMIB or AGP guarantees than before. There are certain results in the literature on psychology and finance that corroborate this hypothesis. The growth of actively managed mutual funds in combination with their often relatively high advisory fees (Freeman and Brown, 2001) shows that customers on financial markets are evidently willing to accept financial advice and even pay for it. Borgsen et al. (2011) show that such advice is more often accepted for insurance policies and Riester contracts than for other financial products. The often complicated structure of variable annuities might increase this tendency even further (Gino and Moore, 2007). Nevertheless, we do not have the data to test this hypothesis. It thus exists as a potential avenue of further research. As long as no empirical evidence exists, it has to be regarded as mere speculation. It is also unclear whether the effect observed here will be persistent over time. Our argument would suggest that this is not the case, because the attention of the individuals will most likely fade from the issue of unisex contracts and the observed effect might disappear. This would be in line with the results of Saito (2006).

6 Conclusion

The question how unisex tariffs impact the demand for insurance has often been answered on a theoretical basis alone (Rothschild, 2011). The empirical analysis of unisex tariffs has in general proven to be complicated because the regulatory implementation has affected all policies at the same time such that a control group is often times not available (e.g. Pope and Sydnor, 2011). Our dataset offers a natural experiment setting which alleviates at least part of the problem commonly associated with empirical analyses in this area.

Our results are not in line with our theoretical predictions. Instead of reacting towards the change in prize for the different contract features, individuals exhibit an overall decreasing demand for contract features in variable annuity products. This effect also persists if the contract features in question become cheaper for the policyholders. While this result could be explained through a sample selection effect, tests for such an effect do not show any evidence for it. We hypothesize that the selective attention in the media might have brought on a wrong perception of unisex priced annuity products and that the consumer reactions might stem from this. However, an empirical test for such an effect will have to be provided by future research.

There are two major recommendations to be drawn from our results. Firstly, at least in the demand for guarantees within variable annuities, rate regulation does not seem to lead to the demand shift necessary for it to cause adverse selection. This would explain the result by Saito (2006). In what sense this result can be applied to the demand for annuities in general or other insurance markets is a subject open to further research. Adverse selection is generally considered to exist in the market for annuities (Mitchell and McCarthy, 2002; Finkelstein and Poterba, 2004). The fact that rate regulation does not worsen adverse selection in this market indicates that rate regulation should also not lead to adverse selection in markets in which asymmetric information play a minor role generally (such as life insurance, e.g., Cawley and Philipson, 1999). As such, the welfare consequences of rate regulation due to adverse selection might generally be less severe than theory suggests.

Several papers have speculated on the effect of rate regulation on adverse selection (Dahlby, 1983; Rothschild, 2011). Some preliminary evidence for it seems to exist, as well. Derrig and Tennyson (2011) use aggregated data on automobile insurance claims to show that the rate regulation in Massachusetts which limits the extend to which high risk drivers can be charged higher premiums leads to higher claims in that state than in other states of the U.S. In a

paper considering the trade-off between predictive accuracy and concerns about discrimination, Pope and Sydnor (2011) show that using full information (including gender) in a statistical model leads to a higher predictive accuracy in predicting claims from unemployment insurance than constrained models. This highlights the potential efficiency loss from rate-regulation, but cannot be considered evidence for adverse selection as insurance coverage is mandatory in their dataset. The fact that our paper cannot find any evidence for the demand shifts required for rate regulation to lead to adverse selection in a natural experiment setting shows that the issue needs to be scrutinized further.

Even though we do not find a change in insurance demand as would be implied by economic theory, we nevertheless observe a difference in demand due to the implementation of unisex tariffs. If our hypothesis is correct that this effect stems from the public perception of unisex tariffs post regulation, our results advice policymakers and insurers alike not to highlight the detrimental consequences of such regulation. Emphasizing the positive aspects instead might even increase insurance demand. However, focusing on the negative consequences of rate regulation seems to be a common reaction by the insurance industry. In 2007, for example, the American Academy of Actuaries wrote that banning the use of information obtained from genetic tests in the underwriting process might "ultimately [result in] raising the cost of insurance to everyone" (American Academy of Actuaries, 2007).

A Proof of Proposition 1

Optimal choices of $(c_i^{\overline{\kappa}}, a_i^{\overline{\kappa}})$ are characterized by the first order conditions:[23]

$$\begin{aligned} V_{c_i}^{\overline{\kappa}} =& -\overline{\kappa}(1-p)\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right) IU_1'(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa})) \\ &+\kappa_i p\int_{\tilde{z}>r}\frac{\partial\gamma}{\partial c_i} IU_2(\gamma(c_i)I)\,\mathrm{d}F(z)=0 \qquad (4) \\ V_{a_i}^{\overline{\kappa}} =& -(1-\overline{\kappa})U_1'(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa}))+(1-\kappa_i)U_B'(a)=0, \end{aligned}$$

We can apply the implicit function rule and obtain

$$\begin{pmatrix} \frac{\partial c_i}{\partial\overline{\kappa}} \\ \frac{\partial a_i}{\partial\overline{\kappa}} \end{pmatrix} = -\frac{1}{\det H}\begin{pmatrix} V_{a_ia_i}^{\overline{\kappa}} & -V_{c_ia_i}^{\overline{\kappa}} \\ -V_{c_ia_i}^{\overline{\kappa}} & V_{c_ic_i}^{\overline{\kappa}} \end{pmatrix}\begin{pmatrix} V_{c_i\overline{\kappa}}^{\overline{\kappa}} \\ V_{a_i\overline{\kappa}}^{\overline{\kappa}} \end{pmatrix},$$

with

$$\begin{aligned} V_{c_ic_i}^{\overline{\kappa}} =& \overline{\kappa}^2(1-p)^2\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right)^2 I^2U_1''(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa})) \\ &+\kappa_i p\int_{\tilde{z}>r}(\frac{\partial\gamma}{\partial c_i})^2 I^2U_2(\gamma(c_i)I)\,\mathrm{d}F(z) \\ V_{a_ia_i}^{\overline{\kappa}} =& (1-\overline{\kappa})^2U_1''(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa}))+(1-\kappa_i)U_B''(a) \\ V_{c_ia_i}^{\overline{\kappa}} =& \overline{\kappa}(1-p)\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right) I(1-\overline{\kappa})U_1''(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa})) \\ V_{a_i\overline{\kappa}}^{\overline{\kappa}} =& U_1'(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa})) \\ &+(\overline{\kappa}-1)(a_i-c_i(1-p)\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right) I)U_1''(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa})) \\ V_{c_i\overline{\kappa}}^{\overline{\kappa}} =& -(1-p)\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right) IU_1'(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa})) \\ &-\overline{\kappa}(1-p)\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right) I(a_i-c_i(1-p)\left(r-\int_{\tilde{z}<r} z\,\mathrm{d}F(z)\right) I)U_1''(w-I-\pi_c(c_i,\overline{\kappa})-\pi_a(a_i,\overline{\kappa})) \end{aligned}$$

[23] Analogously to footnote 12 it can be shown that $(c_i^{\overline{\kappa}}, a_i^{\overline{\kappa}})$ maximizes expected utility.

To determine the impact of a change in $\overline{\kappa}$ on c_i, the implicit function rule yields[24]

$$\begin{aligned}
\frac{\partial c_i}{\partial \overline{\kappa}} \det H &= V^{\overline{\kappa}}_{c_i a_i} V^{\overline{\kappa}}_{a_i \overline{\kappa}} - V^{\overline{\kappa}}_{a_i a_i} V^{\overline{\kappa}}_{c_i \overline{\kappa}} \\
&= \overline{\kappa}(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I(1-\overline{\kappa}) U_1' U_1'' \\
&\quad + \overline{\kappa}(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I(1-\overline{\kappa})(\overline{\kappa}-1)(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I)(U_1'')^2 \\
&\quad + (1-\overline{\kappa})^2(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I U_1' U_1'' \\
&\quad + (1-\overline{\kappa})^2\overline{\kappa}(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I)(U_1'')^2 \\
&\quad + (1-\kappa_i)(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I U_1' U_B''(a) \\
&\quad + (1-\kappa_i)\overline{\kappa}(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I) U_1'' U_B''(a) \\
&= (1-\overline{\kappa})(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I U_1' U_1'' + (1-\kappa_i)(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I U_1' U_B''(a) \\
&\quad + (1-\kappa_i)\overline{\kappa}(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I) U_1'' U_B''(a)
\end{aligned}$$

The first two terms of the sum are negative due to risk aversion and a concave bequest utility function, whereas the third term of the sum is negative if and only if $a_i < c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I$, i.e. if

$$I > I^T := \frac{a_i}{c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)}$$

Therefore $I \geq I^T$ is a sufficient condition for $\partial c_i / \partial \overline{\kappa}$ to be negative.

In case of $I < I^T$, it is a sufficient condition for $\partial c_i / \partial \overline{\kappa}$ to be negative that

$$\begin{aligned}
&(1-\kappa_i)(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I U_1' U_B''(a) \\
&< -(1-\kappa_i)\overline{\kappa}(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I) U_1'' U_B''(a)
\end{aligned}$$

This holds if and only if

$$ARA_{U_1} < \theta_c := \frac{1}{\overline{\kappa}(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I)}. \tag{5}$$

Note that θ_c is positive, as we are in the case of $I < I^T$.

[24] We compress notation by omitting the arguments in U_1, i.e. we write U_1 for $U_1(w - I - \pi_c(c_i, \overline{\kappa}) - \pi_a(a_i, \overline{\kappa}))$.

The implicit function rule furthermore yields

$$\begin{aligned}
\frac{\partial a_i}{\partial \overline{\kappa}} \det H &= V^{\overline{\kappa}}_{c_i a_i} V^{\overline{\kappa}}_{c_i \overline{\kappa}} - V^{\overline{\kappa}}_{c_i c_i} V^{\overline{\kappa}}_{a_i \overline{\kappa}} \\
&= -\overline{\kappa}(1-p)^2 \left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)^2 I^2 (1-\overline{\kappa}) U_1' U_1'' \\
&\quad - \overline{\kappa}^2 (1-p)^2 \left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)^2 I^2 (1-\overline{\kappa})(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I) U_1'' U_1'' \\
&\quad - \overline{\kappa}^2 (1-p)^2 \left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)^2 I^2 U_1' U_1'' \\
&\quad - \overline{\kappa}^2 (1-p)^2 \left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)^2 I^2 (\overline{\kappa}-1)(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I)(U_1'')^2 \\
&\quad - \kappa_i p \int_{\tilde{z}>r} (\frac{\partial \gamma}{\partial c_i})^2 I^2 U_2(\gamma(c_i) I) \,\mathrm{d}F(z) U_1' \\
&\quad - \kappa_i p (\overline{\kappa}-1)(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I) \int_{\tilde{z}>r} (\frac{\partial \gamma}{\partial c_i})^2 I^2 U_2(\gamma(c_i) I) \,\mathrm{d}F(z) U_1'' \\
&= -\overline{\kappa}(1-p)^2 \left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)^2 I^2 U_1' U_1'' - \kappa_i p \int_{\tilde{z}>r} (\frac{\partial \gamma}{\partial c_i})^2 I^2 U_2(\gamma(c_i) I) \,\mathrm{d}F(z) U_1' \\
&\quad - \kappa_i p (\overline{\kappa}-1)(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I) \int_{\tilde{z}>r} (\frac{\partial \gamma}{\partial c_i})^2 I^2 U_2(\gamma(c_i) I) \,\mathrm{d}F(z) U_1''
\end{aligned}$$

The first two terms of the sum are positive due to risk aversion, whereas the third term of the sum is positive if and only if $a_i > c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I$, i.e. if

$$I < I^T := \frac{a_i}{c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right)}$$

Therefore $I \leq I^T$ is a sufficient condition for $\partial a_i / \partial \overline{\kappa}$ to be positive.

In case of $I > I^T$, it is a sufficient condition for $\partial a_i / \partial \overline{\kappa}$ to be positive that

$$\begin{aligned}
&- \kappa_i p \int_{\tilde{z}>r} (\frac{\partial \gamma}{\partial c_i})^2 I^2 U_2(\gamma(c_i) I) \,\mathrm{d}F(z) U_1' \\
&< \kappa_i p (\overline{\kappa}-1)(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I) \int_{\tilde{z}>r} (\frac{\partial \gamma}{\partial c_i})^2 I^2 U_2(\gamma(c_i) I) \,\mathrm{d}F(z) U_1''
\end{aligned}$$

This holds if and only if

$$ARA_{U_1} < \theta_a := \frac{1}{(\overline{\kappa}-1)(a_i - c_i(1-p)\left(r - \int_{\tilde{z}<r} z \,\mathrm{d}F(z)\right) I)}. \tag{6}$$

Note that θ_a is positive, as we are in the case of $I > I^T$. This concludes the proof.

B Discussion of the results of Proposition 1

This section discussed the results of Proposition 1 in a more technical way. As shown above, the results depend on whether we are in the case of $I > I^T$ or the case of $I < I^T$, with

$$I^T := \frac{a_i}{c_i(1-p)\left(r - \int_{\tilde{z}<r} z \, \mathrm{d}F(z)\right)}.$$

I^T gives some indication the ratio of annuity guarantee risk and portfolio downside risk. In case of $I > I^T$, i.e. $c_i(1-p)\left(r - \int_{\tilde{z}<r} z \, \mathrm{d}F(z)\right) I$ larger than a_i, the GMIB risk can be seen as the predominat risk. In case of $I < I^T$, the annuity guarantee would be the predominant risk. As described above, the comparative statics $\partial c_i/\partial\overline{\kappa}$ and $\partial a_i/\partial\overline{\kappa}$ comprise two effects, a direct effect and an indirect substituition effect. For the predominant risk the direct effect always prevails. However, for the other risk the substitution effect might outweigh the direct effect.

An example refering to the fist part of Proposition 1 will make this point clearer. In this case I is larger than the threshold I^T, i.e. the GMIB is the predominant risk. Therefore the result for the choice of c_i is universally valid, as the direct effect prevails. On the other hand, the choice of a_i may be influenced by the substitution effect. Women have a lower GMIB premium in unisex tariffs. Even if they adjust their demand for downside protection upwards, i.e. choose a higher c_f and therefore increase the premium $\pi_c(c_f, \overline{\kappa})$, they might still pay a lower GMIB premium compared to the bisex tariff. This can be the result of a favourable $\overline{\kappa}$ regarding the predominant risk, the GMIB protection, as both premiums, π_c and π_a, have to be paid in the first period. If this is the case, women have to adjust their period-1 consumption to smooth consumption over the two periods. As the only mode to transfer wealth between the period is the VA, this can only be achieved by also adjusting a_f upwards. The higher premium $\pi_a(a_f, \overline{\kappa})$ then compensates the discount on $\pi_c(c_f, \overline{\kappa})$.

Next we want to analyze the thresholds θ_c and θ_a, which determine when the consumption smoothing motive is large enough that the substitution effect prevails. To stick to the example just given, we describe θ_a, i.e. condition (6). The argumentation for θ_c is analogous. As both premiums are paid in the first period, it seems natural that the threshold depends on the individual's risk aversion in the first period. If $ARA_{U_1} < \theta_a$, the consumption smoothing motive is not strong enougt and the direct effect prevails, i.e. a_i increases in $\overline{\kappa}$.

Condition (6) is rather complex. It depends on the guaranteed return r, expectations about the capital market p and $\tilde{z}$, the investment I and the optimal choice $(c_i^{\overline{\kappa}}, a_i^{\overline{\kappa}})$, which in turn depends on the aforementioned parameters. To gain intuition on the threshold, we can look at the case of CARA utility, for which the left-hand side of (6) is constant. We find that an increase in the guaranteed return r decreases the threshold θ_a so that it becomes less likely that (6) is fulfilled. This is plausible, since an increase in r increases the GMIB premium and thus strengthens the substitution effect. The same holds true for an increase in the investment I. In contrast, an increase in the expected performance of the risky asset, i.e. an increase in p or in $\int_{\tilde{z}<r} z \, \mathrm{d}F(z)$ decreases the GMIB premium and thus weakens the substitution effect. In these cases, the threshold θ_a increases so that it becomes more likely that (6) is fulfilled.

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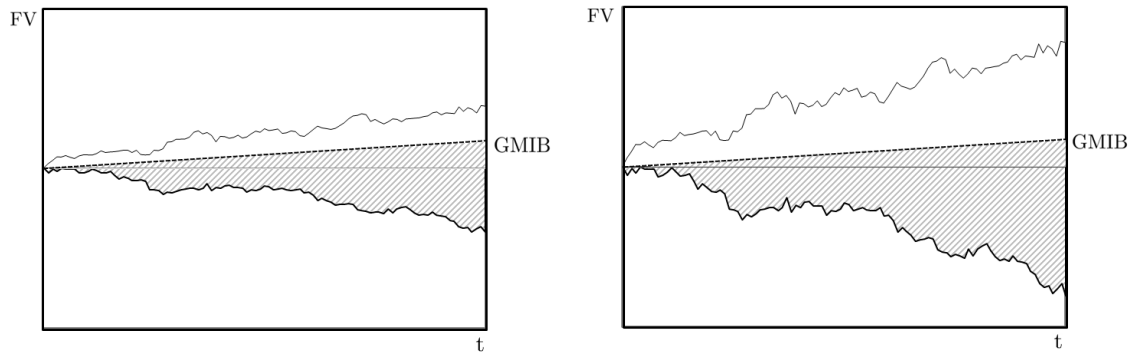
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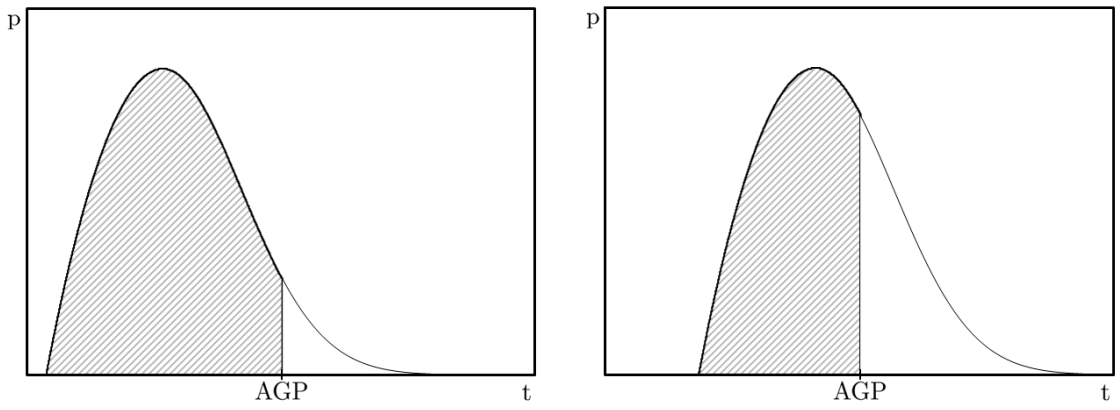
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(a) Low risk portfolio (b) High risk portfolio

Figure 1: Difference in the amount of insurance coverage through the GMIB clause in the annuity contracts for different risk levels. The horizontal axes indicate the time horizon of the investment.



(a) Low life expectancy (b) High life expectancy

Figure 2: Difference in the amount of insurance coverage through the AGP clause in the annuity contracts for different life expectancies.

Table 1: Variable description

Variable	Value	Description
Dependent variable		
Risk	discrete	share of risky asset in portfolio (30%; 50%; 70%)
AGP	continuous	length of guaranteed annuity payment
Independent variables		
$t_{\geq 21.12.2012}$	dummy	1, if commencement date after regulation took effect
treat	dummy	1, if contract is in the treatment group
female	dummy	1, if insured person is female
age	continuous	age at commencement date
duration	continuous	period between commencement and maturity
ln(premium)	continuous	logarithm of the savings premium
invoice	dummy	1, if payment on invoice
year	continuous	year of contract signing

The table reports all descriptive variables used in one or more of the analyses reported in this study. Variables are reported as continuous even when they are only quasi-continuous with a large number of categories such as the guarantee period which is an integer variable.

Table 2: Treatment and control group

	Treatment (non-Riester)	Control (Riester)	$\boldsymbol{\Sigma}$
$<$ *Dec* 21^{st} 2012	4,275	10,671	14,946
$\geq$ *Dec* 21^{st} 2012	1,649	2,169	3,818
$\boldsymbol{\Sigma}$	5,924	12,840	18,764

The table reports the respective sample sizes of the treatment and control group before and after December 21^{st} 2012.

Table 3: Preliminary analysis of risky portfolio share

	Treatment (non-Riester)			Control (Riester)			$\boldsymbol{\Sigma}$
	female	male	$\boldsymbol{\Sigma}$	female	male	$\boldsymbol{\Sigma}$	
$<$ *Dec* 21^{st} 2012	43.87%	44.24%	44.12%	35.61%	37.97%	36.73%	38.84%
$\geq$ *Dec* 21^{st} 2012	43.14%	42.62%	42.87%	38.60%	40.31%	39.41%	40.90%
$\boldsymbol{\Sigma}$	43.61%	43.86%	43.76%	36.12%	38.36%	37.18%	39.26%

This preliminary analysis reports the mean share of risky assets in the portfolios of policyholders in the treatment group and in the control group by gender before and after December 21^{st} 2012.

Table 4: Preliminary analysis of the annuity guarantee period

	Treatment (non-Riester)			Control (Riester)			$\boldsymbol{\Sigma}$
	female	male	$\boldsymbol{\Sigma}$	female	male	$\boldsymbol{\Sigma}$	
$<$ *Dec* 21^{st} 2012	15.07	15.59	15.41	14.36	14.76	14.55	14.80
$\geq$ *Dec* 21^{st} 2012	15.68	15.46	15.57	15.32	15.57	15.43	15.49
$\boldsymbol{\Sigma}$	15.29	15.56	15.45	14.52	14.90	14.70	14.94

This preliminary analysis reports the mean annuity guarantee period in the portfolios of policyholders in the treatment group and in the control group before and after December 21^{st} 2012.

Table 5: Interaction Effects

	Men		Women	
	Control	Treatment	Control	Treatment
$<$ 21.12.2012	β_0	$\beta_0 + \beta_2$	$\beta_0 + \beta_4$	$\beta_0 + \beta_2$ $+\beta_4 + \beta_6$
$\geq$ 21.12.2012	$\beta_0 + \beta_1$	$\beta_0 + \beta_1$ $+\beta_2 + \beta_3$	$\beta_0 + \beta_1$ $+\beta_4 + \beta_5$	$\beta_0 + \beta_1 + \beta_2 + \beta_3$ $+\beta_4 + \beta_5 + \beta_6 + \beta_7$

The table reports how our empirical model predicts the mean share of risky assets in the portfolios and the average annuity guarantee period for the different groups under scrutiny.

Table 6: Estimation results

Estimation	(1)		(2)		(3)		(4)	
Variable	risk	agp	risk	agp	risk	agp	risk	agp
$t_{\geq 21.12.2012}$	0.0268***	0.882***	0.0177**	0.598	0.0235***	0.804***	0.0141	0.525
	(0.00305)	(0.144)	(0.00897)	(0.425)	(0.00458)	(0.211)	(0.00950)	(0.452)
$treat$	0.0738***	0.859***	0.0309***	-0.235*	0.0627***	0.827***	0.0261***	-0.186
	(0.00248)	(0.100)	(0.00331)	(0.138)	(0.00325)	(0.133)	(0.00391)	(0.162)
$t_{\geq 21.12.2012} \times treat$	-0.0392***	-0.727***	-0.0318***	-0.584***	-0.0396***	-0.931***	-0.0339***	-0.829***
	(0.00511)	(0.217)	(0.00531)	(0.221)	(0.00717)	(0.303)	(0.00729)	(0.305)
$female$			-0.00338*	-0.154*	-0.0235***	-0.397***	-0.00797***	-0.174
			(0.00193)	(0.0913)	(0.00234)	(0.123)	(0.00234)	(0.126)
$t_{\geq 21.12.2012} \times female$					0.00639	0.149	0.00309	0.0605
					(0.00612)	(0.289)	(0.00592)	(0.288)
$treat \times female$					0.0199***	-0.117	0.0118**	-0.144
					(0.00514)	(0.208)	(0.00514)	(0.205)
$t_{\geq 21.12.2012} \times treat \times female$					0.00244	0.585	0.00292	0.564
					(0.0103)	(0.439)	(0.0103)	(0.434)
Controls			✓	✓			✓	✓
N	18,764	18,764	18,764	18,764	18,764	18,764	18,764	18,764
adjusted R-squared	0.057	0.005	0.112	0.029	0.062	0.006	0.113	0.029

The table reports the results of the difference in difference estimation indicated in equations (1) and (2). Estimation (1) reports the difference in difference estimation without differentiating between men and women. Estimation (2) adds the vector of control variables. Estimation (3) includes the differentiation between genders but without control variables and estimation (4) includes the full specification with control variables. *, **, and *** indicate significance at the 10%, 5% and 1% level, respectively.

Table 7: Robustness tests for our choice of specification

Estimation	(4)		(5)		(6)	(7)	(8)		(9)	
Variable	risk	agp	risk	agp	risk	risk	risk	agp	risk	agp
$t_{\geq 21.12.2012}$	0.0141	0.525	0.00302	0.309	0.0898	0.0331	-0.0115	-0.0294	0.0177*	0.598
	(0.0095)	(0.452)	(0.0105)	(0.499)	(0.0878)	(0.0943)	(0.0117)	(0.615)	(0.00995)	(0.469)
$treat$	0.0261***	-0.186	0.0138*	-0.672**	0.0899**	0.106*	0.0353***	-2.014***	0.0123***	1.397***
	(0.00391)	(0.162)	(0.00760)	(0.338)	(0.0420)	(0.0632)	(0.00567)	(0.208)	(0.00466)	(0.197)
$t_{\geq 21.12.2012} \times treat$	-0.0339***	-0.829***	-0.0185*	-0.678*	-0.223***	-0.151*	-0.0401***	-0.752*	-0.0300***	-0.682**
	(0.00729)	(0.305)	(0.00946)	(0.412)	(0.0673)	(0.0802)	(0.0105)	(0.421)	(0.00848)	(0.336)
$female$	-0.00797***	-0.174	-0.0221***	-0.431	-0.114***	-0.218***	-0.00667***	-0.201	-0.00675***	-0.189
	(0.00234)	(0.126)	(0.00720)	(0.348)	(0.0265)	(0.0693)	(0.00234)	(0.127)	(0.00234)	(0.126)
$t_{\geq 21.12.2012} \times female$	0.00309	0.0605	0.0136	0.417	0.0916	0.130	0.00264	0.0571	0.00291	0.0674
	(0.00592)	(0.288)	(0.00902)	(0.434)	(0.0585)	(0.0864)	(0.00592)	(0.288)	(0.00591)	(0.288)
$treat \times female$	0.0118**	-0.144	0.0229**	-0.0749	0.163***	0.225**	-0.000946	-0.547*	0.0168***	-0.0785
	(0.00514)	(0.205)	(0.0116)	(0.512)	(0.0453)	(0.0991)	(0.00840)	(0.307)	(0.00604)	(0.224)
$t_{\geq 21.12.2012} \times treat \times female$	0.00292	0.564	-0.00595	0.442	-0.0360	-0.0602	0.0131	1.188*	-0.000703	0.222
	(0.0103)	(0.434)	(0.0146)	(0.640)	(0.0895)	(0.126)	(0.0157)	(0.619)	(0.0120)	(0.481)
Controls	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
N	18,764	18,764	6,476	6,476	18,764	18,764	14,720	14,720	16,884	16,884
adjusted/pseudo R-squared	0.113	0.029	0.056	0.028	0.079	0.034	0.112	0.030	0.121	0.032

The table reports several different specifications for the purpose of demonstrating the robustness of our results. Estimation (5) estimates the equations (1) and (2) using only data from the last contract generation. Estimations (6) and (7) estimate equation (1) as an ordered probit estimation using the full data and only the last contract generation, respectively. Estimations (8) and (9) report the results when the contract group is limited to only the regular contracts (estimation (8)) and only the occupational pension contracts (estimation (9)). *, **, and *** indicate significance at the 10%, 5% and 1% level, respectively. Estimation (4) is reported for ease of comparison.

Table 8: Share of females in treatment and control group

	Treatment			Control	$\boldsymbol{\Sigma}$
	total	regular	occupational		
$<$ *Dec* 21^{st} 2012	34.34%	34.10%	34.45%	52.61%	47.38%
$\geq$ *Dec* 21^{st} 2012	47.67%	46.06%	48.53%	52.91%	50.65%
$\boldsymbol{\Sigma}$	38.06%	37.73%	38.22 %	52.66%	48.05%

The table shows the share of female policyholders with commencement date before and after December 21^{st} 2012. The shares are relative to the total population of each group as indicated by the column headings.

Table 9: Robustness test for endogenous treatment choice

Estimation	(4)		(10)	
Variable	risk	agp	risk	agp
$t_{\geq 21.12.2012}$	0.00792	0.525	0.0315***	0.748
	(0.00947)	(0.449)	(0.0102)	(0.485)
$treat$	0.0146***	-0.185	0.142***	4.296***
	(0.00529)	(0.251)	(0.00841)	(0.393)
$t_{\geq 21.12.2012} \times treat$	-0.0223***	-0.825**	-0.0622***	-1.405***
	(0.00773)	(0.367)	(0.0113)	(0.533)
$female$	-0.00795***	-0.173	-0.00986***	-0.0186
	(0.00250)	(0.119)	(0.00284)	(0.135)
$t_{\geq 21.12.2012} \times female$	0.00306	0.0609	0.00746	-0.0833
	(0.00591)	(0.280)	(0.00760)	(0.362)
$treat \times female$	0.0105**	-0.119	0.0290***	-0.413
	(0.00478)	(0.227)	(0.00700)	(0.329)
$t_{\geq 21.12.2012} \times treat \times female$	0.00427	0.540	-0.0292*	0.333
	(0.00950)	(0.450)	(0.0152)	(0.719)
Controls	✓	✓	✓	✓
Endogenous treatment choice			✓	✓
N	18,764	18,764	18,011	18,011
adjusted R-squared	0.113	0.029		

The table reports the results of estimations accounting for a possible endogenous choice between the treatment group and the control group. *, **, and *** indicate significance at the 10%, 5% and 1% level, respectively. Estimation (4) is reported for ease of comparison.